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The Reliability of Health Risk Appraisals: A Field Trial of Four Instruments

KEVIN W. SMITH, MA, SONJA M. MCKINLAY, PHD, AND JOHN B. MCKINLAY, PHD

Abstract: A field trial was conducted to evaluate the reliability of four widely used health risk appraisals (HRAs). A total of 338 randomly selected adults aged 25 to 65 years residing in the Boston metropolitan area completed an HRA on two occasions seven to 12 weeks apart. Test-retest reliability was assessed by comparing self-reported baseline risk scores to follow-up scores for heart attack risk and selected coronary heart disease risk factors. Respondents generally gave consistent reports for family history, cigarette smoking, and relative weight (test-retest $r > .75$), but self-reported scores for physiologic status (blood pressure and cholesterol) and lifestyle (diet, physical activity, and stress) were much less stable. Coefficients for heart attack risk and appraised age ranged from .43 to .87 for the four HRAs. The reliabilities of two self-scored instruments were greatly reduced by mathematical errors made by respondents when computing their heart attack risk scores. These results were not affected by the length of the follow-up period (seven to 12 weeks). (*Am J Public Health* 1989; 79:1603–1607.)

Introduction

Reliability refers to the degree to which measurements are affected by random error. Random error effects may be assessed by the stability of the measures or scores produced by an instrument from one time period to another. The higher the correlation between scores measured at two points in time under identical conditions, the more reliable the instrument.[1]

For health risk appraisal (HRA) instruments designed to provide estimates of mortality or morbidity risk for various diseases, reliability is an important issue for several reasons. First, reliability is related to the validity of an HRA. Random errors in self-reported health status and physiologic measures will decrease the accuracy of an HRA's estimate of disease risk. Second, HRAs are frequently employed to evaluate the effectiveness of health promotion programs. If HRA scores are unstable during the pre- to post-program periods, changes attributable to the intervention may be indistinguishable from unreliability in the instrument itself. Moreover, respondents may be inclined to dismiss HRA estimates if they find that their risk scores fluctuate dramatically from one time period to another.

In recent years, several studies have raised questions about the reliability of HRA data. Best and Milsum found that participants in a smoking cessation program reported many improbable changes in their body frame size, medical history, and family history after six months.[2] Sacks and his associates found that only 15 percent of the subjects in a clinical trial gave consistent responses to all HRA items at both baseline and follow-up interviews.[3] On the other hand, some studies have reported test-retest reliability coefficients exceeding .7 for selected risk factors and overall risk estimates, especially for follow-up periods of short duration (three days to one month).[4–7]

Several methodological limitations associated with these studies make it difficult to generalize about HRA reliability.

- First, some investigators have reported only the percentage of respondents choosing the same response category for a particular item at both points in time. This is inadequate for assessing the degree of reliability because it does not take into account the differences between the risk scores assigned to the categories. In many instruments, selecting an adjacent category for a particular item produces little or no change in the risk score for that item; in other cases, the difference may be substantial.
- Second, much of the previous work has been based on limited samples of self-selected volunteers. In some instances, fewer than 30 subjects were involved in the analyses.[5] It is not clear whether the results from these studies are representative of the general population.
- A third problem common to all test-retest analyses is that reliability may be confounded with actual change. Scores for the same individual may differ from one administration to another because other behavioral changes have occurred in the intervening period.
- Finally, the effect of the length of the follow-up period on reliability coefficients for HRAs has yet to be established.

Recognizing these shortcomings, several recent reviews of the status of risk appraisal methods have called for more comprehensive investigations of HRA reliability.[8–10]

In response to these issues, a field trial was designed to assess the reliability and validity of four HRAs. For each HRA, this paper examines:

- the proportion of respondents choosing the same risk factor categories at both administrations,
- test-retest reliability coefficients for individual risk factors and total risk scores,
- the effect of behavioral changes occurring during the follow-up interval on reliability coefficients, and
- the influence of the length of the follow-up period on changes in risk scores.

Methods

The field trial was conducted from January to November 1987. Subjects were identified through random digit dialing of 15 telephone exchanges in the Boston, Massachusetts metropolitan area. Five exchanges were drawn from each of three strata comprising calling areas with below average, average, and higher than average household incomes. One adult was selected at random from each household and screened for eligibility. The study was restricted to persons between the ages of 25 and 65 at last birthday who did not

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have a history of coronary heart disease, diabetes, or hypertension, so that the sample would reflect the typical target population for HRAs that contain coronary heart disease components. An introductory letter explaining the purpose of the study was sent to each respondent who provided an address at the time of the screening.

Eligible cases were then contacted by a trained field technician to arrange a home visit. During the visit, the technician administered a standardized personal interview and one of the four HRAs evaluated by the project. Field technicians were not permitted to help respondents complete HRA items or compute risk scores. Subjects were reinterviewed seven to 12 weeks after the initial visit. Fifty-five percent of the eligible cases (the reliability group) were asked to complete the same HRA at follow-up. The follow-up interval was designed to be long enough to minimize memory effects, but short enough so that major changes in lifestyle and behavior would be unlikely to occur. The analyses presented in this paper are restricted to field trial participants completing the same HRA at both baseline and follow-up.

Four widely-used HRAs were chosen to represent the major types of instruments that are currently available:[11]

1) Health Risk Appraisal (Centers for Disease Control)
2) The Heart Test (Arizona Heart Institute)
3) RISKO (American Heart Association)
4) Determine Your Medical Age (Blue Cross/Blue Shield, NY)

Each instrument requires a respondent to select categories that best describe personal habits, physiologic status, and medical history. Point values or weights associated with these categories are then combined to produce an overall estimate of risk. In these four HRAs, overall risk is summarized by either a heart attack risk scale, where higher values represent greater risk, or by appraised age. Appraised age is the age of adults with the same all-causes mortality risk as a particular individual. The items and range of risk scores for each HRA are shown in the Appendix.

Reliability analyses were conducted by comparing the baseline scores for individual risk items, heart attack risk, and appraised age with the values chosen by the same respondent at follow-up. Consistent with psychometric practice, Pearson correlations between the baseline and follow-up risk scores served as reliability coefficients.[12] These correlations are approximately the same as weighted Kappas with weights equal to squared deviations between baseline and follow-up risk scores.[13,14] For four items, the influence of other changes occurring during the follow-up period was assessed by computing test-retest correlations only for those who said, in another part of the interview, that their behavior had not changed since baseline. For relative weight, diet, and physical activity, these correlations were restricted to those reporting that their behavior at follow-up was "about the same" as it had been at baseline. Analyses for smoking were based only on those who reported smoking the same number of cigarettes at both interviews. Finally, follow-up risk scores were regressed on baseline scores and the number of days that had elapsed between the dates of the baseline and follow-up were computed to evaluate the impact of the length of the follow-up interval on reliability.

In the absence of established minimum standards for test-retest coefficients, correlations of .8 or greater were adopted a priori as a criterion level in this study. This level is equivalent to the informal standard for internal consistency used in many social science applications.[12] Sample sizes for each of the four HRA groups in the field trial were designed to have sufficient power to detect a difference between the standard of .8 and a correlation of .65 or less, which was regarded as unacceptably small for reliable items.[15]

Results

Baseline interviews were completed by 732 of 1,276 eligible adults (57.4 percent) contacted for the field trial. Table 1 compares the demographic characteristics reported during the telephone screening by reliability group participants, other participants in the field trial, and eligible subjects who declined to participate. All three groups were similar with respect to these measures; a logistic regression analysis of response rates indicated that the probability of obtaining a completed baseline interview was not affected by these characteristics (likelihood ratio statistic = 5.70, df = 6). Of the 403 reliability group subjects interviewed at baseline, 338 (84 percent) consented to a follow-up interview and completed HRAs at both administrations (n = 83 to 87 persons for each instrument).

Response Category Comparisons for Risk Items

Table 2 compares the percentage of respondents who chose the same category for eight major risk items at both baseline and follow-up. The marginal distributions of these items were similar at each administration. Participants were generally very consistent in their reports of family history and smoking. Choices for relative weight were consistent at least 70 percent of the time. The lowest agreement rates occurred for lifestyle factors such as diet, stress, and physical activity, which may change even over short periods of time.

The consistency of blood pressure and cholesterol reports are more difficult to evaluate. These reports were the same at both administrations at least 87 percent of the time for two of the three self-scored instruments. Determine Your Medical Age, however, merely specifies a cut-off level so that only two choices were available to respondents. Similarly, the Arizona Heart Test specifies only three categories for blood pressure: normal, borderline, or high. The agreement rates for RISKO were below 70 percent, but this instrument offered more category options than the others. Like most of the Geller/Gesner mortality risk HRAs, the Centers for Disease Control (CDC) instrument requests actual blood pressure readings and cholesterol levels. The majority of participants left these items blank because they did not know

TABLE 1—Demographic Characteristics of Field Trial Participants and Nonparticipants[*]

Characteristics	Reliability Group Participants (N = 403)	Other Field Trial Participants (N = 329)	Non-Participants (n = 544)
Age (years)	38.8 (10.6)	40.0 (10.8)	40.2 (11.1)
Percent female	55.8 (49.7)	59.9 (49.1)	60.5 (48.9)
Percent Black/Hispanic	6.7 (25.1)	10.7 (30.9)	8.2 (27.4)
Income (thousands)[a]	41.1 (10.4)	41.2 (11.1)	40.5 (10.1)
Self-reported health status[b]	2.06 (.94)	1.92 (.86)	2.06 (.91)
Worry about health[c]	2.54 (.83)	2.54 (.85)	2.52 (.84)

[*] Mean and (standard deviation)
[a] Mean household income in subject's telephone exchange
[b] 5-point scale; 1 = excellent; 5 = poor
[c] 4-point scale; 1 = not at all, 4 = most of the time

TABLE 2—Percent of Respondents Choosing Same Risk Item Category at Baseline and Follow-up by HRA

Risk Items	Arizona Heart Institute	CDC	Determine Your Medical Age	RISKO
Family History	91.9	96.4	86.2	—
Smoking	82.6	91.6	83.9	85.7
Relative Weight	79.8	78.0	71.3	80.8
Physical Activity	64.3	73.5	65.5	—
Blood Pressure	94.6	NR	87.4	58.4
Cholesterol	—	NR	94.3	64.8
Stress	65.6	—	—	—
Diet	72.3	—	55.2	—

NR = value not reported by majority of respondents

the appropriate values. For this reason, no results for these two risk factors are presented for CDC in Table 2.

Test-Retest Correlations for Risk Items

The proportion of respondents selecting the same response categories at two points in time is, at best, a crude indicator of risk factor reliability because it fails to account for the degree of risk involved.[16] For the self-scored instruments, different response categories sometimes have the same risk values, so that selecting different categories may not result in any change in the risk score for that item. To provide a more sensitive measure of reliability, test-retest correlations were computed for each HRA on the basis of the scores associated with each risk factor category. The CDC instrument differs from the others in that risk weights rather than point values are assigned to item categories.

Table 3 shows the test-retest correlations for the risk scores associated with the eight HRA items. Correlations between baseline and follow-up risk scores for family history, smoking, and relative weight exceeded .75 for all applicable instruments. The remaining five items all had estimated reliability coefficients that were significantly lower than the criterion level of .8 for at least one HRA. In spite of differences in question wording and risk scores, the reliability estimates for a specific item were usually quite similar among the four HRAs.

Effects of Behavioral Change on Reliability

Discrepancies between pre- and posttest reports do not always mean that information has been reported unreliably because behavior may have changed since the baseline interview. Test-retest correlations were therefore computed for selected items only for those respondents who claimed their behavior had not changed during the follow-up interval. These items were weight (relative weight), exercise level (physical activity), sodium, fat, and caloric intake (diet), and smoking risk (only those who gave the same number of cigarettes per day at both interviews).

Results for these respondents are summarized in Table 4. Compared to the coefficients for all cases (Table 3), behavioral changes had the greatest impact on reports of dietary habits. The estimated reliabilities increased slightly (.04 on average) for smoking and relative weight, but were virtually identical for physical activity risk scores. Therefore, with the possible exception of dietary practices, it would appear that behavioral changes did not greatly influence the reliabilities of these risk items.

Test-Retest Correlations for Heart Attack Risk and Appraised Age

In the process of collecting data during the field trial, it quickly became apparent that many respondents who received self-scored HRAs had difficulty computing these risk scores. Heart attack risk scores and appraised age values were therefore checked for accuracy by project staff. Eighty-seven percent of those completing the Arizona Heart Institute HRA were able to calculate the correct score compared to 75 percent of the RISKO respondents and only 42 percent of those completing Determine Your Medical Age. These rates reflect the mathematical skills required by each instrument—the Arizona Heart Institute HRA involves adding point values, RISKO requires both addition and subtraction, while respondents must add, subtract and divide scores to determine their "medical" age. Error rates were also influenced by the educational attainment of respondents. Those without four-year college degrees were twice as likely as college graduates to make computational mistakes on both the RISKO (22.5 percent compared to 9.0 percent) and Determine Your Medical Age (61.4 percent compared to 30.0 percent) HRAs.

Test-retest reliabilities for heart attack risk scores and appraised age were calculated separately for the risk scores actually reported by respondents as well as score totals corrected for mathematical mistakes. The coefficients for reported risk scores were .87 for the Arizona Test HRA, .84 for the CDC HRA, and .74 for Determine Your Medical Age. The correlation between baseline and follow-up scores reported by RISKO respondents (r = .43), on the other hand, was significantly lower than the criterion level of .8. For this instrument, even minor addition or subtraction errors can produce misleading estimates of whether a person is above or below the average risk for heart disease. When these errors were corrected, RISKO's estimated reliability increased to .76. Correcting computational errors also produced a marked improvement in the reliability of Determine Your Medical

TABLE 3—Risk Item Test-retest Correlations and 95 Percent Confidence Intervals by HRA

	Arizona Heart Institute		CDC		Determine Your Medical Age		RISKO	
Risk Items	*r*	95% CI	r	95% CI	r	95% CI	r	95% CI
Family History	.90	(.84–.93)	.75	(.64–.83)	.89	(.83–.93)	—	—
Smoking	.89	(.84–.93)	.95	(.93–.97)	.82	(.73–.88)	.89	(.83–.93)
Relative Weight	.84	(.76–.89)	.96	(.94–.97)	.78	(.68–.85)	.78	(.67–.85)
Physical Activity	.60	(.45–.73)	.65	(.50–.76)	.70	(.58–.80)	—	—
Blood Pressure	.87	(.78–.92)	NR	—	.45	(.26–.60)	.55	(.36–.69)
Cholesterol	—	—	NR	—	.65	(.51–.76)	.58	(.40–.72)
Stress	.62	(.47–.74)	—	—	—	—	—	—
Diet	.52	(.34–.66)	—	—	.57	(.41–.70)	—	—

TABLE 4—Risk Item Test-retest Correlations and 95 Percent Confidence Intervals for Cases Reporting No Change in Behavior during Follow-up Interval

	Arizona Heart Institute		CDC		Determine Your Medical Age		RISKO		Mean Change in r*
Risk Items	r	95% CI	r	95% CI	r	95% CI	r	95% CI	
Smoking	.95	(.92–.97)	.97	(.95–.98)	.86	(.78–.91)	.95	(.93–.97)	+.04
Relative Weight	.88	(.80–.93)	.96	(.94–.98)	.76	(.61–.85)	.91	(.84–.95)	+.04
Physical Activity	.67	(.49–.79)	.62	(.41–.77)	.66	(.46–.80)	—	—	.00
Diet	.78	(.63–.87)	—	—	.63	(.43–.77)	—	—	+.16

* = mean difference from correlation in total HRA sample

Age (from r = .74 to r = .99) but not the Arizona Heart Institute HRA, for which mathematical mistakes were much less common. While computational mistakes are not an issue with computerized Geller/Gesner method HRAs like the CDC's, these instruments typically insert risk weights of 1.0 for any unknown risk factors. This practice artificially inflates reliability coefficients.

Length of Follow-up Period

The reliability of most measurements decreases over time. In this study, the interval between baseline and follow-up interviews ranged from seven to 12 weeks (49 to 84 days). The final set of analyses was designed to determine whether the reliability of respondents' self-reports declined as the length of the follow-up period increased. To address this issue, follow-up risk scores were regressed on baseline scores and the number of days that had elapsed between the baseline and follow-up interviews. The regression results (Table 5) indicate that the length of the follow-up interval had little impact on the follow-up scores for any of the risk measures. However, this interval was comparatively short and had limited variability among respondents.

TABLE 5—Multiple Regression of Follow-up Risk Score on Length of Follow-up Interval* Adjusted for Baseline Risk Score: Unstandardized regression coefficient and (standard error)

Risk Score	Arizona Heart Institute	CDC	Determine Your Medical Age	RISKO
Heart Attack Risk	.044	28.2	—	.046
	(.047)	(21.7)	—	(.093)
Appraised Age	—	.040	−.192	—
	—	(.027)	(.166)	—
Family History	−.001	.001	−.014	—
	(.012)	(.001)	(.023)	—
Smoking	.028	.002	.045	−.009
	(.019)	(.002)	(.070)	(.007)
Relative Weight	.000	.001	.058	−.013
	(.010)	(.001)	(.054)	(.011)
Physical Activity	.027	.004	.149	—
	(.016)	(.007)	(.087)	—
Blood Pressure	−.021	NR	−.015	−.011
	(.012)	—	(.045)	(.012)
Cholesterol	—	NR	.005	−.016
	—	—	(.019)	(.010)
Stress	.019	—	—	—
	(.023)	—	—	—
Diet	.034	—	−.022	—
	(.020)	—	(.056)	—

*Length of follow-up interval = days elapsed between baseline and follow-up interviews
NR = value not reported by majority of respondents

Discussion

In this field trial, respondents generally gave consistent reports for three risk items: family history, cigarette smoking, and relative weight. Self-reports of physiological status (blood pressure and cholesterol) and lifestyle (diet, physical activity, and stress), on the other hand, were much less consistent between the baseline and follow-up interviews. The consistency of physical activity reports may have been diminished by seasonal variations in exercise patterns.

There were some important differences among the instruments with respect to the reliabilities of heart attack risk scores and appraised age. The test-retest coefficients for the Arizona Heart Institute and Centers for Disease Control HRAs both exceeded .84, but were only .74 and .43 for Determine Your Medical Age and RISKO, respectively. The correlations for the latter two instruments were reduced by mathematical errors made by respondents when computing their risk scores. Reliability coefficients for the CDC HRA, on the other hand, may be overstated due to the practice of substituting risk weights of 1.0 for blood pressure and cholesterol whenever the actual values are unknown.

Inconsistencies between baseline and follow-up values were more likely to be attributable to unreliable reporting on the part of respondents rather than to actual changes in risk factor status. Changes in risk scores were not affected by variations in the length of the follow-up period. However, testing reactivity could have lowered reliability levels. Completing an HRA for the first time can motivate respondents to obtain more accurate information about family history, blood pressure, or cholesterol, so that follow-up responses will then appear to be inconsistent.

The results of the field trial indicate that the reliability of HRA risk scores can vary greatly from one instrument to another. These findings have several implications for health educators using HRAs. First, computation errors may severely reduce the reliability of self-scored instruments. Computerized HRAs may therefore be the most desirable type unless manual calculations can be routinely checked for errors. Second, additional work needs to be done to improve the measurement of specific risk factors, particularly physical activity levels, dietary practices, and physiologic status. The reliability of blood pressure and cholesterol levels can be increased by measuring these values during physical examinations rather than relying on self-reports. Finally, as the reliability of an HRA declines, it becomes increasingly difficult to distinguish changes in risk status from random reporting errors. This is particularly important for investigators interested in assessing the effects of health promotion interventions because unreliability makes intervention effects more difficult to detect. Therefore, only HRAs for which reliability can be demonstrated should be considered for evaluations of the effectiveness of these interventions.

APPENDIX

Risk Scores Ranges by HRA

	Arizona Heart Institute (self-scored)	Centers for Disease Control (computer-scored)	Determine Your Medical Age (self-scored)	RISKO (self-scored)
Heart Attack Risk	0–50	2–21,969	—	0–27
Appraised Age	—	18.4–65.1	12.0–70.6	—
Family History	0–5	.9–1.2	−30–15	—
Smoking	0–10	.8–2.1	−6–24	−2–4
Relative Weight	0–3	.8–1.6	−6–24	−2–3
Physical Activity	0–3	.6–2.5	−12–12	—
Blood Pressure	0–6	.4–1.1	0–12	−5–6
Cholesterol	—	.5–1.6	0–6	−3–3
Stress	0–6	—	—	—
Diet	0–6	—	−6–18	—

The Centers for Disease Control (CDC) HRA, which requires computer processing, provides probability estimates for the 12 leading causes of death as well as appraised age based on the Geller/Gesner method.[17]

The Arizona Heart Institute Test and RISKO are self-scored instruments focusing on heart attack risk.

In the Determine Your Medical Age HRA, appraised age is calculated by adjusting an individual's chronological age for medical history and lifestyle factors.

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Call for Papers:
Risk Management in Rural Health Care

The American Society for Healthcare Risk Management has issued a call for papers to be considered for its Fall 1990 issue of *Perspectives in Healthcare Risk Management*. Emphasis should be on the issues and problems inherent in rural health care risk management, as well as on suggestions for how to deal with them effectively. Topics and approaches include but are not limited to: credentialing issues, peer review, conflict of interest, patient transfer issues, swing-beds, risk pooling, and quality assurance/risk management integration.

Authors should submit a one-page abstract and detailed outline of an original article on any area of risk management in rural health care delivery settings. *Abstracts and outlines must be received no later than March 5, 1990*. Authors of those abstracts selected for further consideration will be asked to submit their completed papers (not to exceed 20 typewritten, double-spaced pages) by June 4, 1990. Manuscripts will be accepted with the understanding that their content is unpublished and not being submitted for publication elsewhere. Send abstracts to:

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American Society for Healthcare Risk Management
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Accuracy of Perceptions of Heart Attack Risk: What Influences Perceptions and Can They Be Changed?

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Abstract: Using perceived risk of a heart attack, we examined the relative importance of perceived risk factors and sociodemographic variables on subjects' perception of heart attack risk, the relationship between perceived and objectively measured heart attack risk, and the effect of health risk appraisal (HRA) feedback on risk perceptions. Data derive from a random sample of 732 Greater Boston, Massachusetts area men and women ages 25–65 years, who participated in a field trial of health risk appraisal instruments. At baseline and approximately two months later, all respondents completed a questionnaire assessing their own health-related behavior, risk factors, and perception of heart attack risk. At baseline, respondents also completed one of four HRA instruments. Physiologic measures of cholesterol, blood pressure, and weight were taken at either baseline or follow-up visit. Results showed that respondents used established risk factors in estimating overall risk; that compared to objective measures of risk, a high percentage of respondents displayed an optimistic bias; and that HRA feedback had some effect on perception of heart attack risk for those at high risk. Implications of these results for health promotion are discussed. (*Am J Public Health* 1989; 17:1608–1612.)

Introduction

A person's beliefs about his or her perceived susceptibility to a condition or disease figure prominently in models of health behavior such as the Health Belief Model,[1,2] the Theory of Reasoned Action,[3] Subjective Expected Utility Theory,[4,5] and Protection Motivation Theory.[6,7] These models assume that perceived susceptibility (also referred to as vulnerability or risk) is essential in motivating behavior.

Previous research, however, has shown that people tend to underestimate their own risk of developing certain conditions or diseases[8–17]—or they have what Weinstein[13] refers to as an "optimistic bias." This underestimation has important implications for health promotion and behavior change programs. If people do not perceive themselves as vulnerable to a disease or condition, they are less likely to adopt recommended behaviors.[1,18] It is therefore important to understand what determines a person's perception of risk and how these perceptions can be made more realistic.

Most of the research on perceived susceptibility uses comparative risk judgments to examine how people compare their risk for a particular condition or disease to that of their peers.[8–10,13–16] This method, however, does not address the accuracy of an individual's risk perception. In an attempt to examine this issue, Weinstein conducted two studies[15,16] in which an individual's comparative risk judgments were correlated with his or her relevant risk factors. Weinstein found that these associations were often weak or nonexistent. However, the measures of risk factors lacked precision and were based only on self-report. In addition, each risk factor was separately correlated with perceived risk, rather than combined to form an overall estimate of risk. One purpose of this study is to use epidemiologically based models to examine the relation between perceived risk and objectively measured risk for heart attacks.

Optimistic biases may have several sources. They may be due to cognitive errors (e.g., lack of information or experience with a condition) or motivational needs (e.g., self-esteem enhancement or defensive denial).[16] To the extent that they result from cognitive errors, individualized feedback about risk may overcome these biases. Health risk appraisal instruments (HRAs) are designed to estimate a person's mortality or morbidity risk for various diseases based on characteristics such as medical history, blood pressure, and smoking habits.[19] One of the goals of HRAs is to provide risk-related feedback that will enable a person to reduce risk in areas amenable to modification.[20] Another purpose of this study is to examine the effect of individualized risk feedback on risk perception.

Methods

Procedure and Sample

The data used in this study were collected during a field trial of health risk appraisal instruments conducted by the Cambridge Research Center of the American Institutes for Research in 1987. Detailed descriptions of the procedure and sample characteristics are reported in the article in this issue by Smith, McKinlay, and McKinlay.[21] Briefly, respondents were identified through random digit dialing of residential households in 15 neighborhoods and towns in the Boston metropolitan area. Eligibility was restricted to adults between the ages of 25 and 65 who did not have a history of coronary heart disease, diabetes, or hypertension. Following a telephone screening, baseline interviews were conducted by trained field technicians in respondents' homes. After completion of the baseline interview, respondents were randomly assigned to complete one of four risk appraisal (HRA) instruments (CDC, Arizona Heart Test, RISKO, and Determine Your Medical Age), or were assigned to a "no HRA" condition. (See Smith, *et al*,[21] for details of the HRAs). Field technicians were not permitted to help respondents complete HRA items or compute risk scores. Follow-up home visits were conducted seven to 12 weeks after the baseline to monitor changes in health status, health practices, and medical care utilization. During one of the two home visits (randomly determined), technicians measured blood pressure, height and weight, and drew a venous blood sample for cholesterol analysis.

Measures

Perceived risk of heart attack or stroke was assessed by the question: "Compared with persons of your own age and

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sex, how would you rate your risk of having a heart attack or stroke within the next 10 years?" Respondents answered on a 5-point scale ranging from "much lower than average" to "much higher than average."

Respondent characteristics that might be related to perceived risk were divided into sociodemographic characteristics, self-perceptions of health, and cardiovascular risk factors. Sociodemographic factors included education, race, gender, and age (although gender and age can also be considered risk factors). Self-perception of health status was measured on a 5-point scale ranging from excellent to poor.

Cardiovascular risk factors were measured as follows. Family history of heart disease was based on having a biological parent die of heart disease or stroke before age 65. Smoking was based on self-report of how many cigarettes per day the person currently smoked. Self-report of smoking is generally thought to be reliable, especially for respondents who are not in a smoking reduction program.[22] All respondents who smoked at least one cigarette per day were classified as smokers. The Harvard Alumni Activity Scale,[23] which measures recreational physical activity in terms of kilocalories (kcal) expenditure per week, provided the measure of exercise. Stress was determined by the number of stressful events (e.g., death of partner or spouse, divorce, loss of job, major change in health of family member), as derived from the Holmes and Rahe Scale[24] that were reported by the respondent as occurring within the past two months. Self-reports of weight, cholesterol, and blood pressure as compared to average were also obtained.

Results

Perceived Heart Attack Risk

Almost 56 percent of the respondents rated their risk as lower than average, 29 percent rated it about average, and only 13 percent rated their risk as higher than average. With over half the respondents rating their risk as lower than average, these results support previous findings that people tend to rate their own risk as lower than that of their peers—that people have an optimistic bias. Although respondents were healthier than their peers because those with hypertension, diabetes, and other chronic diseases were excluded from the study, this optimistic bias is confirmed when compared to objective risk (see below).

To examine the relative importance of perceived risk factor status, sociodemographic factors, and perceived health in people's perception of heart attack risk, a multiple regression analysis was performed in which perceived risk of a heart attack reported at baseline was the dependent variable. These results show associations between perceived heart attack risk and the following risk indicators: number of cigarettes smoked, death of a parent due to heart disease; and self-assessments of weight, blood pressure and cholesterol compared to average levels (see Table 1). Neither amount of exercise nor number of recent stressful events was related to perceptions. These results suggest that people do make use of cardiovascular risk indicators in estimating overall risk. These risk factors accounted for 24 percent of the variance in perceived risk. Self-reported health and education were also positively related to perceived risk and, combined with the risk factors already included, accounted for 30 percent of the variance. The correlation between self-reported health and risk is partly a function of the way in which self-reported health was scored (i.e., a 5-point scale). The Pearson correlations between these variables and perceived risk show a pattern similar to the multiple regression, suggesting that multicollinearity is not a problem in this analysis.

TABLE 1—Results of Multiple Regression on Variables Related to Perceived Heart Attack Risk at Baseline

Variable	Unstandardized Regression Coefficient	S.E.
Self-reported risk factors		
Cigarettes smoked per day	.02	.00
Cholesterol relative to average	.24	.05
Weight relative to average	.23	.05
Death of parent due to CHD	.38	.10
Blood pressure relative to average	.19	.06
Exercise	−.03	.03
Number of recent stressful events	−.02	.06
Perceived Health	.34	.05
Sociodemographic Factors		
Education	.08	.03
Age	−.01	.00
Minority	−.22	.13
Female	.12	.08

Education was positively related to perceived risk in the regression model, yet had a negative Pearson correlation with perceived risk. Because Pearson correlations revealed that education was related to self-perceived health, exercise, smoking, and blood pressure, and because we thought education might modify these effects, we also performed a multiple regression that included the interactions between education and these variables. The only significant result was an interaction between education and smoking, such that those with more education who smoked perceived their risk to be greater.

Perceived and Objective Heart Attack Risk

The risk equation from the new Carter Center Healthier People HRA was used to obtain an objective measure of coronary heart disease risk.[25] This logistic risk model builds on the results of the Risk Factor Update Project but takes into account more recent epidemiological data from the Framingham Heart Study. The variables included in the logistic equation to predict mortality from myocardial infarction are: sex, total cholesterol, age, number of cigarettes smoked per day, systolic blood pressure, relative weight, HDL cholesterol, and presence of diabetes. Family history and exercise are not included in the model because it is thought that their effects operate through other risk indicators in the model (cholesterol, blood pressure, and weight).[25]

Because respondents were instructed to answer the risk perception question in comparison to other people of the same age and sex, we adjusted the objective measure of risk for age and sex. Each respondent received a risk ratio which was based on the logistic estimate from the model divided by the average risk for that person's age, sex, and race. A ratio of 1.0 meant a person was at average risk for his or her age, sex, and race.

Respondents were then classified as to whether they were at above average, average, or below average risk. A ratio below .75 was considered below average risk, .75 to 1.25 average risk, and over 1.25 above average. In order to create a symmetric distribution, these ratios were transformed onto a log scale, in which the cut-offs became ln(.75) and ln(1.33). These ratios represent equal "risk distances" from the null value of 0.0 on the log scale [$\ln(1.33) = -\ln(.75) = .29$]. Using this distribution, 39 percent of the respondents were classified as below average risk, 24 percent as having average risk, and 36 percent as above average risk.

Perceived risk was then compared with the objectively measured risk categories. In Table 2, the rows indicate the percentage of respondents in each risk perception category, while the columns indicate the percentage of respondents in each objectively measured category. Summing over cells, a total of 42 percent of the respondents underestimated their risk (15% + 17% + 10%), while only 18 percent (11% + 3% + 4%) overestimated risk, and 40 percent (25% + 9% + 6%) were accurate. Based on these results, when using an objective measure of risk, this analysis also found evidence of an optimistic bias—a high percentage of respondents underestimated their risk.

We next performed two logistic regressions: one to predict the probability that risk was overestimated (pessimists), and the other to predict underestimation of risk (optimists). Variables included as predictors were the same variables listed in Table 1 that were used to predict perceived heart attack risk. The variables associated with pessimism were death of parent due to heart disease (O.R. = 2.72, 95% CI = 1.52, 4.87), self-assessed health status (O.R. = 1.80, 95% CI = 1.25, 2.58), and age in years (O.R. = 1.06, 95% CI = 1.02, 1.08). Thus, pessimists were more likely to be in worse health, have had a parent die of heart disease, and be older. The result for death of a parent may reflect the fact that people give weight to this risk factor while the Carter Center model does not include it. The variables that predicted optimism were age (O.R. = .90, 95% CI = .87, .92) and education (O.R. = .80, 95% CI = .66, .98). Those who were younger and had less education were more likely to be optimists.

Effect of Health Risk Feedback

Table 3 shows the effect of feedback on *change* in perceived risk. In this table, the columns indicate the type of feedback received, while the rows indicate whether perceived risk increased, did not change, or decreased from baseline to follow-up. The numbers in parentheses are the percentage of respondents who did not change their perceptions and whose perceptions accurately reflected their HRA feedback. In reading this table, it is important to keep in mind that those who received feedback that they were below average risk should have either been accurate or decreased their perceived risk, while those who received feedback that they were above average risk should have either been accurate or increased their perceived risk.

As seen in Table 3, the majority of respondents did not change their risk perceptions. Those who changed the most received feedback that they were above average risk. They were more likely to increase their perceived risk (21.8 percent of those receiving above average feedback). This group, however, was also the least accurate in their risk perceptions. Those who received feedback that they were at below average or average risk, or did not receive any HRA feedback were as likely to increase their perceived risk as to decrease it. It should be pointed out that because respondents tended to underestimate their risk at baseline, there were more respondents in the "above average" feedback category with the potential to change. While HRA feedback had its greatest effect on those at above average risk, 12 percent of those with above average risk changed their perceptions in the opposite direction of the feedback.

TABLE 2—Relationship between Perceived and Estimated Risk as Measured by the Carter Center Model

	Risk Ratio			
Perceived Risk[a]	Below %	Average %	Above %	Total
Lower than Average	25	15	17	57
Average	11	9	10	30
Higher than Average	3	4	6	13
Total	39	28	33	100
n	(226)	(138)	(205)	(569)

[a] As compared to persons of same age and sex.

TABLE 3—Relationship between HRA Feedback and Change in Perceived Risk

	HRA Feedback[a]			
Change in Perceived Risk	No Feedback	Below Average Risk	Average Risk	Above Average Risk
Increase	16.9	10.4	14.9	21.8
No Change	69.7	78.7	74.5	66.3
(accurate)		(60.8)	(26.6)	(16.8)
Decrease	13.5	11.0	10.6	11.9
	100%	100%	100%	100%
n	89	347	94	101

[a] Scored 1 = below average, 2 = average, 3 = above average.

To further explore whether receipt of HRA feedback altered risk perception, we looked at follow-up risk perception as a function of baseline risk perception, the same variables examined in the baseline risk perception analysis (this time using follow-up responses) plus HRA feedback. A multiple regression was performed in which follow-up risk perception was regressed on baseline risk, follow-up perceptions of risk factors, sociodemographic factors, perceived health, and HRA feedback. Feedback was measured by a scale relating HRA risk category to the respondent's initial perception. The respondent's initial perception was taken into account because feedback should only have an effect when it differs from the person's original perception. The feedback scale ranged from 1 to 5, where 1 meant the HRA risk category was much lower than the individual's perception, and 5 meant the HRA risk category was much higher than the respondent's perception.

These results, shown in Table 4, reveal only feedback from the Arizona Heart Institute (AHI) HRA had a significant impact on heart attack risk perception. Compared to the other three HRAs, the feedback from this HRA was presented in a concrete fashion directly related to heart disease risk. Respondents were clearly told that they were above average, average, or below average risk for developing heart disease. Thus, respondents may be more likely to understand and use the feedback from the AHI.

Risk Perception and Behavior

Although it was not a primary purpose of the study, because data on respondents' risk-related behavior were available, it was possible to examine the relationship between change in perceived risk from baseline to follow-up and corresponding behavior change. As previously mentioned, models of health behavior often assume a relationship between perceived susceptibility and behavior.[1–6] It follows from this assumption that increases in perceived risk should be related to risk-reducing behaviors. To address this issue,

TABLE 4—Results of Multiple Regressions on Variables Related to Perceived Heart Attack Risk at Follow-Up

Variables	Unstandardized Regression Coefficient	S.E.
Baseline Risk Perception	.56	.04
Risk Factors		
Cholesterol relative to average	−.12	.04
Cigarettes smoked per day	.01	.002
Weight relative to average	.11	.04
Death of parent due to CHD	.11	.08
Blood pressure relative to average	.09	.05
Exercise	−.02	.03
Number of recent stressful events	−.05	.04
Self-reported health	.12	.04
Sociodemographic Factors		
Education	.03	.03
Minority	−.04	.10
Female	−.03	.06
Age	.003	.003
Health Risk Appraisal Feedback[a]		
CDC	.01	.07
RISKO	.07	.07
Arizona Heart Institute	.26	.10
Medical Age	.09	.07

[a] Scored relative to baseline perception of risk.

we looked at behavior change from baseline to follow-up as a function of change in risk perception. The behaviors we analyzed were: smoking, exercise, weight; self-reports of reduction in salt, fat, and calories; and recent blood pressure or cholesterol screening. Results showed that those who increased their perceived risk of a heart attack were no more likely to change their behavior than those whose risk perception either decreased or remained the same.*

Discussion

Our results showed that people do use established epidemiologic risk factors in estimating their overall risk of heart attack, but that additional factors also influence a person's risk perception. One such variable is self-perceived health status: those people who perceive their overall health as worse are more likely to rate their risk as higher. This finding is consistent with other research in which people who were sick estimated their risk for other illnesses as higher.**

Using an objective measure of individual risk, a high percentage of respondents exhibited an optimistic bias, providing additional support for previous studies.[9,10,13–16] Those who were younger and less educated tended to be more optimistic, consistent with suggestions that younger people feel less vulnerable to conditions or diseases.[16] The finding that family history of heart disease is associated with pessimism suggests that people weigh this factor heavily and is consistent with Becker and Levine's[8] study of high risk siblings of patients with coronary heart disease that found less optimistic bias than seen in other studies.

The effect of health risk feedback on risk perceptions was greatest for those who were above average in risk: twice the percentage of respondents who were above average risk *increased* their perceived risk. These findings suggest that HRA feedback can effectively change the perceptions of those at high risk. It is important, however, to add a cautionary note: 12 percent of those who received high risk feedback actually *decreased* their perceived risk and this above average group was the *least* accurate to begin with. Furthermore, the results of the multiple regression showed that only feedback from the AHT was significantly associated with change in risk perception. Clearly, feedback did not always have consistent effects on risk perception and these effects differed by HRA.

With respect to various explanations of the optimistic bias, to the extent that feedback had the desired effect on accuracy, the cognitive explanation of this bias is supported. However, it is also clear that information did not have the desired effect for the majority of respondents.

In the present study we did not find a relation between increases in perceived risk and behavior change. Although some models of health behavior predict a relation between perceived susceptibility and behavior,[1–6] support for this association is strongest for disease prevention behaviors that do not require modification of health habits.[26] Our behavioral measures consisted of changes in personal health habits such as smoking, diet, exercise, and weight. Models of health behavior also suggest that factors other than cognitions (e.g., social support, social norms and self-efficacy) may be important in health habit modification.[3,26,27] It should be pointed out, however, that this analysis was based on limited statistical power as only 87 respondents increased their perceived risk.

Two important aspects of this study need to be pointed out. One is that the logistic model of actual risk is based on mortality from heart attack, while respondents were asked about their chances of *having* a heart attack or stroke. It is possible that they may have responded differently if asked about their chances of dying from a heart attack or stroke, although we feel it is unlikely the respondents made this distinction. The other important point has to do with HRA feedback. Respondents were only given the feedback provided by the HRA and did not receive any additional counseling or education, as is often done with HRAs. In some cases, respondents may not have understood the feedback.*** Thus, HRAs may not be effective for use in mass screenings, but may have a greater effect on risk perception when feedback is easier to understand, or when they are given in conjunction with a health education program, a conclusion also drawn by Schoenbach, *et al*,[28] in a review of the effectiveness of HRAs.

In conclusion, this study suggests that people do make use of established risk factors in estimating cardiovascular risk; that compared to objective measures of risk, a high percentage of people display an optimistic bias; and that risk feedback has some effect on risk perception for those at high risk, but such feedback is not optimally effective.

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*Data available on request to author.

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NCHSR Grants $4 Million for Patient Outcome Research

The National Center for Health Services Research and Health Care Technology Assessment (NCHSR) has awarded new grants totaling nearly $4 million in first-year costs in a series that will measure the effectiveness of alternative medical procedures in treating specific health conditions. Total spending on the nationwide DHHS program will be $5.9 million this fiscal year, with $52 million proposed for FY 1990.

The new grants will support research for treatment of conditions that affect a large number of Americans every year—prostate gland enlargement, heart attack, cataracts, and lower back pain—for which millions of Medicare beneficiaries are treated every year. The research projects—expected to require about five years to complete—will assess comparative effects of alternative therapies on patient survival, quality of life, costs, and other factors. Research teams will consist of practicing clinicians and other experts.

Grants to support patient outcome research assessment teams were awarded to:

Richard Deyo, MD, University of Washington ($896,049), nonsurgical interventions for low back pain and different types of lumbar surgery;

Barbara J. McNeill, MD, Harvard Medical School ($900,000), for a national data base for outcomes from myocardial infarction treatment;

Earl P. Steinberg, MD, Johns Hopkins University ($899,986), assess cataract procedures;

John E. Wennberg, MD, Dartmouth Medical School ($933,535), transurethral resection, open prostate surgery, and nonsurgical interventions for enlargement of prostate gland.

Planning grants of $50,000 each awarded to seven research teams included:

Fred Featherstone, MD, American Academy of Orthopedic Surgery, hip replacement/disease; *Alvan Feinstein*, MD, Yale University, use of colonoscopy; *Emmett B. Keeler*, PhD, Rand Corporation, treating peripheral vascular diseases; *David B. Matchar*, MD, Duke University, prevention and management of stroke; *Lawrence W. McMahon, Jr.*, MD, University of Michigan, use rates and appropriateness of cholecystectomy, and screening/treatment for colon cancer; *David B. Pryor*, MD, coronary artery disease; and *Sanford Schwartz*, MD, University of Pennsylvania, gallstone disease.

For further information, consult *Research Activities* (special release), from Publications and Information Branch, NCHSR, 18–12 Parklawn Building, Rockville, MD 20857; Tel: 301/443-4100.

4 T

The Prevention of Silicosis and Prediction of Its Future Prevalence in China

JIEZHI LOU, MD, AND CHEN ZHOU, MD, MPH

Abstract: We studied the effects of preventive and therapeutic interventions on the 18-year cumulative incidence of silicosis of 26,603 dust-exposed workers in seven Chinese mines and industrial plants. Cumulative silicosis incidence decreased from 36.1 percent in workers employed before 1950 to 1.5 percent in workers employed after 1960. From the 1950s to 1970s, eight-year cumulative incidence of tuberculosis decreased from 54.7 percent to 16.7 percent and case fatality of silicosis patients dropped from 53.9 percent to 18.3 percent. From 1950s to 1980s, the average age at the detection of silicosis increased from 41.3 to 52.7 and the average survival time of silicosis patients prolonged from 2.0 to 12.2 years. Workers over age 40 who began exposure to dust before 1960 will be the main source of new silicosis patients in future. Most expected new cases of silicosis will occur within the next 15 years. The silicosis population will remain unchanged for the next 20 years and will decrease dramatically after 25 years. There will be few silicosis patients in 30 years and new cases of silicosis will be difficult to find in 45 years. These results show that the preventive and therapeutic actions against silicosis in China have been effective. (*Am J Public Health* 1989; 79:1613–1616.)

Introduction

Silicosis is a major occupational lung disease which remains a problem in industrialized and developing countries.[1,2] Since the 1950s, special attention has been paid to the prevention of silicosis in China. A regulation to control the silica dust hazard in industrial establishments and mines was published in 1956,[3] and a regulation concerning the prevention and treatment of silicosis was published in 1957.[4] Experiences with dust control were summarized and applied nationwide in 1958. During the past 30 years, many industrial companies have implemented these regulations which include monitoring dust concentrations in the work environment, applying wet mining and engineering control methods, screening for silicosis in dust-exposed workers, and improving the care of silicosis patients. The evaluation of the effects of these preventive and therapeutic efforts on the change in prevalence of silicosis and predicting its future prevalence are important for future preventive programs.[5–7] We, therefore, conducted an investigation on the prevalence and development of silicosis in seven industrial companies of different types. Based on these data, we predicted the future trend of silicosis in these establishments.

Methods

Subjects and Their Industrial Environments

The subjects of this investigation had been exposed to dust for at least three years and had physical examination, occupational history records, and chest X-rays available for the study. The 26,603 dust-exposed workers came from two metallic mines (one lead mine, one molybdenum mine), two coal mines (coal mine 1, and coal mine 2), a fireproof raw material workshop in a steel plant, a foundry in a large machinery plant, and a raw material workshop in a glass factory. These seven industrial companies were selected as representative of different industrial environments with silica dust hazards and also because their records were complete.

Preventive and Therapeutic Interventions

From 1955 to 1957, the two metallic mines and two raw material workshops initiated comprehensive dust control programs focusing on wet operation. The other three establishments took similar actions in 1962 or 1963. Although there was no systematic dust monitoring data, the dust concentration were very high according to a few random measurements. After applying wet mining and engineering control methods, the average dust concentration decreased sharply (Table 1). Dust was collected on membrane filter and measured by gravimetric methods. Free silica content was measured by chemical-gravimetric method.

Silicosis patients were systematically screened and identified from the raw material workshops in 1950 and 1951 respectively. Similar screening in the two metallic mines—coal mine 2, and the foundry—began in 1955–56 and 1958–59, respectively, while the first screening for silicosis in coal mine 1 began in 1963. As the numbers of silicosis patients increased, the companies set up special physician-directed organizations to examine dust-exposed workers and to treat the silicosis patients. Dust-exposed workers were systematically screened annually and approximately one-third of the dust-exposed workers were examined by chest X-rays each year. More than 90 percent of the silicosis patients were

TABLE 1—Average Dust Concentrations (total dust) in Seven Industrial Companies before and after Dust Control Regulations

Companies	Date (years)	No. of Samples	Mean (mg/M^3)	SiO$_2$ (%)
Lead	1956	18	340.0	31.5
	1957–81	2038	5.7	
Molybdenum	1956	24	313.1	5.0
	1957–81	43621	4.7	
Refratory	1956	33	287.7	49.2–98.6
	1957–81	4417	3.9	
Cast Steel	1959–62	111	200.5	85.0–97.0
	1963–81	144	29.3	
Glass	1953–56	28	263.5	92.0–98.5
	1957–81	185	4.3	
Coal 1	1960–62	24	211.0	31.0
	1963–81	5646	17.2	
Coal 2	1960–62	25	222.5	23.0
	1963–81	1242	17.8	

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transferred from jobs that exposed them to dust as well as given symptomatic treatment. Some companies also set up a special center to care for simple silicosis patients.

Diagnostic Criteria

The diagnosis of silicosis was performed by special silicosis diagnostic teams in each province or municipality. The criteria of radiographic classification of silicosis was based on the silicosis diagnostic criteria published by the Ministry of Health of People's Republic of China:[4]

> Stage I silicosis: definite evidence of nodular opacities. This demands at least 10 nodular opacities of 1 mm or larger in diameter in two square centimeters in any field of lung.
>
> Stage II silicosis: nodular opacities increase in number and enlarge in size and occupy more than four of the six lung fields.
>
> Stage III silicosis: nodular opacities conglomerate into large opacities of a minimum of 2 cm in diameter.
>
> Silicotuberculosis: tuberculosis developed in any stage of silicosis.

These criteria were set up for silicosis. They are crude and do not consider irregular opacities and profusion in comparison with ILO (International Labour Office) criteria, but they are accurate for silicosis diagnosis and easy to use.

If the subject was dead, the date of death, age at death, and cause of death were recorded. We interviewed the workers or their relatives and co-workers in a few cases when records were incomplete. We also recorded all workers who were no longer available because of non-silicosis related death or relocation.

Analysis

We divided the seven companies into three groups to represent three different industrial environments: metallic mines group (lead mine, molybdenum mine); coal mine group (coal mine 1 and coal mine 2); and plants group (a fireproof raw material workshop, a raw material workshop in a glass manufacturing factory, a foundry). We analyzed the prevalence of silicosis, case fatality rate, and causes of death among these three groups. To obtain the cumulative incidence rate, we divided observed workers into four cohorts according to the time of their first exposure to dust. The number of workers entered in these four cohort were 3,331, 7,108, 7,907 and 8,257, respectively, and the workers were followed up for 18 years, e.g., those with first exposure to dust before 1950 were followed up until 1966. A life table method was used to calculate the 18-year cumulative incidence of silicosis among workers who began exposure to dust before 1950, during 1950–54, 1955–59, and 1960–64.[8] The average annual incidence of silicosis was calculated as follows:

average annual incidence (%) = cumulative incidence (%)/observed years.

We also divided silicosis patients into three cohorts according to the time of diagnosis. Rates of silicosis progression, silicosis mortality and complication with tuberculosis in each eight-year diagnosis period.

To predict the future prevalence of silicosis, we used the average annual incidence rate derived from cumulative incidence of silicosis in dust-exposed workers, life expectancy of healthy dust-exposed workers of different age groups to predict the expected new cases of silicosis among healthy workers and the future trend of silicosis. The life expectancy of healthy dust-exposed workers of different age groups was calculated according to the life expectancy of male residents in Shenyang, China in 1976.[9] The details of this method were described in a previous report.[10]

Results

Silicosis

By the end of 1981, 3,803 silicosis patients were diagnosed (14.3 percent of the 26,603 dust-exposed workers). The rate in workers exposed to dust before 1950 was highest (36.1 percent), rates for workers employed later decreased gradually. For those who started employment in 1960 or later, the rates in the metallic mine group and the plant group decreased to 0.30 and 0.07 percent, respectively; the rates of the coal mine group remained somewhat high (3.0 percent) because of delay in dust suppression (Table 2).

The 18-year cumulative incidence of silicosis was highest in workers who started exposure to dust before 1950 and declined in those whose exposure started later (Table 3).

As the incidence of silicosis declined, the average age at disease occurrence and average duration from starting exposure to development of stage I silicosis (latent period) increased. The average age at which silicosis was diagnosed was 41.3 and average latent period 11.3 years in the 1950s, increasing to 45.7 and 14.5 in the 1960s, 48.7 and 17.0 in the 1970s, and 52.7 and 19.7 in the 1980s, respectively.

Disease Progression, Tuberculosis as a Complication

By the end of 1981, of the 3,803 diagnosed silicosis cases, 997 (26.2 percent) had progressed to a subsequent stage and 1,634 (43.0 percent) had been diagnosed as having tuberculosis. The average percent of progression (progressing from stage I to stage II and III or from stage II to stage III) was 26.2 percent and the average incidence of tuberculosis was 43.0 percent. Among the 1,634 cases of silicosis complicated with tuberculosis, 901 cases (55.1 percent) were identified at the time when silicosis was first diagnosed. The life table method was used to calculate the cumulative incidence of progression and of tuberculosis among patients diagnosed before 1955, in 1955–64 and in 1965–74 (Figures 1 and 2).

Death from Silicosis

By the end of 1981, 1,094 of 3,803 silicosis patients had died, a mortality rate of 28.8 percent. Figure 3 shows the case fatality within equal periods of eight years of the patients diagnosed in different time periods. The case fatality declined over time. Table 4 presents the average age at death and mean duration from silicosis diagnosis to death among three groups during different time periods. The average age at death and mean duration from diagnosis to death increased significantly over time.

TABLE 2—Cases of Silicosis among Workers Employed in Different Time Periods

Industry Group	Date (year)	No. of Workers	No. of Silicosis Cases	Percent
Metallic mine	1950	1106	442	39.96
	1950–	3281	397	12.10
	1955–	1980	84	4.24
	1960–78	3021	9	0.03
Coal mine	1950	1886	595	31.55
	1950–	2874	840	29.23
	1955–	5154	767	14.88
	1960–78	3814	115	3.02
Plant	1950	339	166	48.97
	1950–	953	309	32.42
	1955–	773	78	10.09
	1960–78	1422	1	0.07
Total	1950	3331	1203	36.12
	1950–	7108	1546	21.75
	1955–	7907	929	11.75
	1960–78	8257	125	1.51

TABLE 3—Comparison of Accumulated Silicosis Prevalence Rates in Worker Cohorts Beginning Exposure to Dust in Different Time Periods

Observed year	Time Entry into Cohorts			
	Before 1950	1955–54	1955–59	1960–64
0–	0.00	0.00	0.00	0.00
1–	0.03	0.00	0.01	0.00
2–	0.12	0.00	0.01	0.00
3–	0.18	0.04	0.01	0.00
4–	0.63	0.17	0.02	0.03
5–	0.93	0.41	0.13	0.06
6–	1.12	0.74	0.21	0.13
7–	3.23	1.16	0.47	0.33
8–	5.09	1.57	0.66	0.84
9–	6.79	2.19	0.81	1.42
10–	8.43	2.84	1.41	1.70
11–	9.04	3.69	1.92	1.80
12–	9.42	4.80	2.77	1.87
13–	10.29	6.05	3.46	1.97
14–	12.58	7.19	4.06	2.36
15–	17.28	8.58	4.61	2.64
16–	18.47	9.96	5.51	2.98
17–	20.49	10.95	6.38	3.28
18–	—	—	—	—

TABLE 4—Average Age at Death and Mean Duration from Diagnosis to Death (survival time) among Silicosis Patients in Different Time Periods

Industry Group	Time (year)	No. of Deaths	Age at Death (years)	Survival Time (years)
Metallic mine group	1950–	23	42.6	1.9
	1960–	112	47.4	5.8
	1970–	128	56.2	11.0
	1980–	48	61.2	14.3
Coal mine group	1950–	—	—	—
	1960–	74	51.2	3.0
	1970–	358	53.4	6.7
	1980–	101	56.5	10.1
Plant group	1950–	30	39.5	2.0
	1960–	90	44.7	5.0
	1970–	94	54.2	10.1
	1980–	36	59.3	16.8
Total	1950–	53	40.8	2.0
	1960–	276	47.5	4.8
	1970–	580	54.2	8.2
	1980–	185	58.0	12.2

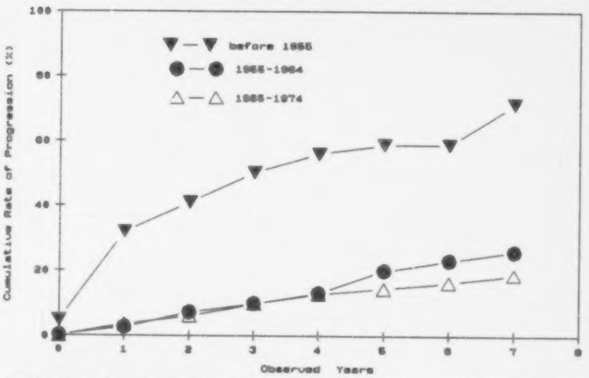

FIGURE 1—Cumulative Rate of Progression of Silicosis among Silicosis Patients Diagnosed at Different Time Periods.

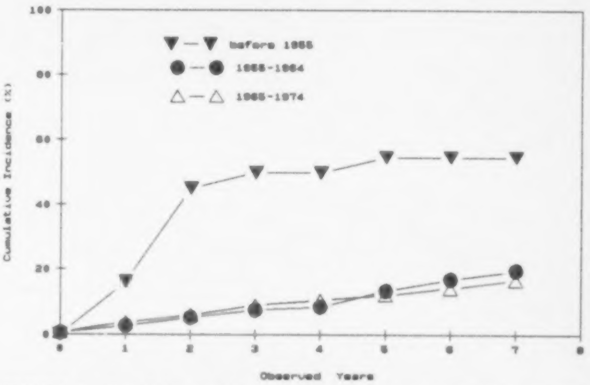

FIGURE 2—Cumulative Incidence of Tuberculosis among Silicosis Patients Diagnosed at Different Time Periods.

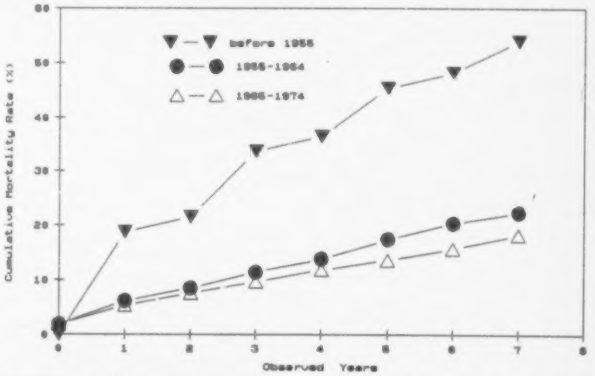

FIGURE 3—Cumulative Mortality Rate among Silicosis Patients Diagnosed at Different Time Periods.

The distribution of specific causes of death of silicosis patients is shown in Table 5. Tuberculosis was a major cause of death (62.1 percent). After improvement in health care, the proportion of death caused by tuberculosis dropped from 98.1 percent in the 1950s to 48.7 percent in the 1980s. With the increase in the mean age at death, the proportion of other pulmonary diseases, cardiovascular diseases, and cancer as causes of death increased from 1.9 percent in 1950s to 44.7 percent in 1980s.

Predicting New Cases of Silicosis

Using annual incidence rates of silicosis in dust-exposed workers, age distribution, and life expectancy of healthy dust-exposed workers, we calculated the expected number of cases of silicosis in the future (Tables 6 and 7). We estimate 464 new cases of silicosis are predicted to develop in the remaining 8,146 healthy workers in the metallic mine group, 1,412 cases are predicted to occur in 10,284 healthy coal miners and 248 cases in 2,520 healthy workers in the plant group. Workers over age 40 who started their exposure to dust before 1960 will be the main source of new cases of silicosis. Most expected cases of silicosis will occur within the next 15 years. The development of new cases of silicosis will be less after 35 years and will not be found in 45 years.

Discussion

During the rapid industrialization of China in the 1950s,

TABLE 5—The Distribution of Specific Causes of Death of Silicosis Patients in Different Time Periods

Time (year)	No. of Deaths	Tuberculosis		Pulmonary Disease		Cancer		Cardiovascular Disease		Other Disease	
		No.	%	No.	%	No.	%	No.	%	No.	%
1950–	53	52	98.1	1	1.9	0	0.0	0	0.0	0	0.0
1960–	276	202	73.2	30	10.9	12	4.3	2	0.7	30	10.9
1970–	580	335	57.8	93	16.0	60	10.3	38	6.6	54	9.3
1980–	185	90	48.7	38	20.5	27	14.6	17	9.2	13	7.0
Total	1094	679	62.1	162	14.9	99	9.0	57	5.2	97	8.8

TABLE 6—Expected Numbers of Silicosis among Healthy Dust-exposed Workers at Different Age Groups

Industry Group	No. of Workers	Expected Numbers of Silicosis at Different Ages						
		25–	30–	40–	50–	60–	70–	Total
Metallic mine	8146	0.7	22.8	63.5	283.7	80.3	13.1	464.1
Coal mine	10284	1.8	208.4	267.8	719.4	214.8	11.8	1424.0
Plant	2520	0.4	25.8	13.3	129.8	37.9	3.9	211.1
Total	20950	2.9	257.0	344.6	1132.9	333.0	28.8	2099.2

TABLE 7—Expected New Cases of Silicosis among Healthy Dust-exposed Workers with First Exposure to Dust at Different Time Periods

Entry Time of Cohort	No. of Workers	Projected New Cases of Silicosis in Future Years						
		<5	5–	15–	25–	35–	45–	Total
<1950	1889	108.8	203.2	75.2	0.7	0.0	0.0	387.9
1950–	5087	177.4	337.1	167.0	2.9	0.0	0.0	684.4
1955–	6376	139.7	276.6	187.9	24.2	0.0	0.0	628.4
1960–	2688	24.1	48.2	43.5	30.9	3.4	0.0	150.1
1965–	4910	32.9	65.9	64.2	57.9	27.4	0.1	248.4
Total	10950	482.9	930.0	538.8	116.6	30.8	0.1	2099.2

dust concentrations in the industrial environment were very high. In the seven establishments we studied, total dust concentrations were between 200 and 300 mg/M^3. The prevalence of silicosis under these high concentrations was about 36 percent, the latent period was 11.3 years, and patients died within two years after diagnosis. But since the Chinese government set up the regulations concerning the control of silica dust and prevention and treatment of silicosis in 1956 and 1957, all of these companies have made special efforts to prevent silicosis. Compared with the 1950s, the cumulative incidence of silicosis, incidence of tuberculosis in silicosis patients, and case fatality rate all declined dramatically. The average age at silicosis occurrence and survival time of silicosis patients increased significantly. These changes strongly demonstrated the effectiveness of preventive and therapeutic efforts in these companies. These silicosis prevention experiences could provide a valuable reference for rapidly industrializing countries.

Our results demonstrated that the prevalence of silicosis in dust-exposed workers decreased with improved working conditions and abatement of the dust concentration. In the coal mines, where dust level had not been completely controlled under the maximum allowable concentration, the prevalence of silicosis in workers employed after 1960 was still relatively higher than in other companies which had lower dust concentrations. Further action to suppress dust in these mines is necessary.

Recently, many researchers paid much attention to predicting epidemiological trends of silicosis incidence.[6,7,10] Since many factors could affect silicosis incidence, it is difficult to estimate the trends accurately. The life table method we used here is a simple but effective approach. With the average annual incidence rate derived from the cumulative incidence of silicosis, the age distribution and life expectancy data of healthy dust-exposed workers, new cases of silicosis could be estimated accurately.

In this study, we found that the workers age 40 and older who were exposed to dust before 1960 will be the major source of new silicosis patients. This result is consistent with those of two investigations reported by Lou, *et al.*[10,11] The future preventive work in these companies should concentrate on early detection of silicosis among these older workers and delay the development of the disease.

Since this investigation was done in 1985 and most new cases of silicosis will develop in the next 15 years, the silicosis population will not decrease within this century. Also, because the survival time of silicosis is between 10 to 15 years, the silicosis patient population will decline sharply after 25 to 30 years. In the years 2020 or 2030 (35 or 45 years later), new cases of silicosis will be very difficult to find. These predictions show that silicosis will remain a problem for the next 40 years, therefore, it is also necessary to improve health care delivery for silicosis patients and develop new anti-silicosis drugs or therapies to prolong their life expectancy.

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3T

Risk of Carotid Endarterectomy in the Elderly

Elliott S. Fisher, MD, MPH, David J. Malenka, MD, Neil A. Solomon, BA, Thomas A. Bubolz, PhD,
Frederick S. Whaley, PhD, and John E. Wennberg, MD, MPH

Abstract: We used the Medicare claims files to describe operative mortality for 2,089 New England residents over the age of 65 who underwent carotid endarterectomy in 1984 and 1985. For patients ages 65 to 69, the risk of death within 30 days of surgery was 1.1 percent, (95% confidence interval = 0.5, 2.1), for those ages 70 to 74, 2.8 percent (1.7, 4.4), for those ages 75 to 79, 3.2 percent (1.8, 5.2), and for those over age 80, 4.7 percent (2.3, 8.5). Nearly 80 percent of patients underwent surgery at hospitals performing 40 or fewer carotid endarterectomies per year on the Medicare population. The adjusted odds ratio for 30 day mortality for patients undergoing surgery in these low-volume hospitals was 2.8 (95% CI = 1.1, 7.2) compared to higher volume hospitals. Although the Medicare claims data provided only limited data about post-operative strokes, analysis of post-operative stroke risk supported these findings. (*Am J Public Health* 1989; 79:1617–1620.)

Introduction

Carotid endarterectomy is commonly performed, costly and controversial. In 1985, more than 100,000 operations were performed in the United States at a cost of $1.2 billion.[1,2] While the operation may reduce the risk of stroke compared to medical management, the currently available scientific evidence is insufficient either to reject or endorse this hypothesis.[2–5] Only two randomized trials on this question have been completed, each with inconclusive results.[6,7] Clinical uncertainly about the appropriate indications for the procedure is reflected in the high degree of regional variation in the utilization rates of carotid endarterectomy.[8,9] For example, residents of Boston, Massachusetts are twice as likely to undergo carotid endarterectomy as residents of New Haven, Connecticut.[10]

The potential benefit of the operation depends primarily on the operative mortality and the risk of stroke with medical or surgical management and high post-operative complication rates virtually eliminate the potential benefit of the operation for all patient subgroups.[3,11,12] Many argue that carotid endarterectomy should be restricted to specialized centers[13,14] or to those with documented operative stroke plus death rates under 5 or 6 percent.[15]

Published case series demonstrate wide variation in complication rates, with stroke plus death rates ranging from 3 percent to over 15 percent.[9,16–19] Three small studies found no significant increased risk in patients over the ages of 75 or 80 and advocated the use of the procedure in elderly patients.[20–22] Low surgical volume has been found to be associated with increased risk for many surgical procedures. Carotid endarterectomy has been specifically examined in two studies, neither of which found an effect of surgical volume on outcome. However, the statistical power of these two studies was limited, as was the spectrum of institutions included in their analyses.[23,24]

We used the Medicare claims files for the six New England states to address three issues: the effect of age on the operative risk of carotid endarterectomy, the association of low surgical volume with increased risk, and the evaluation of the performance of carotid endarterectomy at individual hospitals.

Methods

Data Source

Data for the current study were obtained from three computer files maintained by the Health Care Financing Administration:[25]

- the Medicare Provider Analysis and Review (MEDPAR) file, which contains computerized discharge abstracts for each hospitalization for Medicare enrollees and includes patient and hospital identifiers and data on diagnoses, procedures, and resource utilization;
- the Health Insurance Skeletonized Eligibility Writeoff (HISKEW) file which records whether a patient died and the date of death, regardless of whether that death occurred within a hospital; and
- the Provider of Services file, which describes the relevant characteristics of all Medicare institutional providers.

Selection of the Study Population

We reviewed all 2,171 hospital discharge abstracts submitted for New England resident Medicare beneficiaries 65 years of age and older discharged between April 1, 1984 and June 30, 1985 who underwent carotid endarterectomy (ICD-9-CM procedure code 38.12). After eliminating the 81 patients who received surgery outside New England and one patient with inconsistent coding, 2,089 patients were selected for inclusion in the study. If the endarterectomy was not listed as the primary procedure (n = 83, 3.8 percent of cases), we assigned the date of surgery for the index endarterectomy to the midpoint of the hospital stay. (We repeated all reported analyses using the date of admission as the date of surgery for these 83 cases; none of the analyses were affected.)

Outcomes

Deaths were determined by searching the write-off file. Stroke was ascertained by inference from the diagnostic codes recorded on the discharge abstract for the index admission and readmissions within 30 days of surgery. If patients' diagnosis codes could be consistent only with a post-operative stroke (n = 35), they were coded as having definite post operative strokes. If the coding of the stroke was ambiguous and thus might have represented a pre-operative stroke, the patient was coded as having a possible stroke (n = 161).* Based on these definitions, we calculated the

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*Details of coding available on request to authors.

proportion of patients who experienced the following outcomes within 30 days of surgery: death; death or definite post-operative stroke (the minimum estimate of the stroke plus death risk); death or any possible post-operative stroke, which represents the maximum risk of death plus stroke that can be ascertained from the discharge abstract (some additional strokes might be detected through chart review). Confidence intervals for these proportions were calculated according to the exact method.[26]

Independent Variables

Patients' age, sex ,and race were ascertained from the discharge record. To identify important medical conditions of patients, we used diagnostic codes from both the surgical (index) admission and from any admissions whose discharge date fell within the six months preceding the date of the index surgical procedure, as previously described.[27] If an index admission diagnosis could reasonably be interpreted as a potential complication of the endarterectomy, a patient was coded as having that condition only if it was also present on admissions that preceded the one in which the endarterectomy occurred. Variables evaluated as risk factors for adverse outcomes were: coronary artery disease, pulmonary disease, diabetes, renal disease, congestive heart failure, valvular heart disease, stroke, and the presence of any hospital admission during the previous six months.

Hospital surgical volume was defined based on the number of index carotid endarterectomies performed by each hospital on the Medicare population during the 15-month study period. Institutional size was defined as a continuous variable based on the number of hospital beds. Hospital teaching status was defined as a continuous variable based on the number of housestaff in accredited residency programs per hospital bed.

Analysis

We used multiple logistic regression to test the association between each of the three outcomes and the clinical and institutional predictors described above.[28] Of the clinical variables considered, only coronary artery disease was significantly associated with 30-day mortality after controlling for age, sex, and surgical volume. Both sex and coronary artery disease were included in the reported logistic regression models. Alternative specifications of the model, in which teaching status and hospital size were treated as both dichotomous and ordinal variables, were examined. Regardless of the specification of the model, the results were similar to those presented below. Two-way interactions between surgical volume and age, and age and sex were examined and were not significant.

We used the binomial distribution to test the hypothesis that the observed number of deaths plus definite strokes at an individual hospital exceeded the number that would have been expected had the hospital's outcomes equaled those of the highest volume hospitals.

Results

Operative Mortality

Operative mortality increased markedly with age (Table 1). Those over age 70 faced a progressively increased risk, with those over age 80 facing a more than four-fold increase in the adjusted odds of death compared to those ages 65 to 69. Operative risk was also higher in hospitals with low surgical volume on the Medicare population (Table 1). Patients undergoing carotid endarterectomy in hospitals in each of the three lower volume groups faced an increased risk of death compared to those receiving surgery in hospitals performing more than 40 endarterectomies per year. The increased risk associated with low-volume surgery was consistent across all specifications of the model examined. When a dichotomous variable was substituted for the three levels of lower hospital volume in an otherwise identical model, the adjusted odds ratio for 30-day mortality at hospitals performing 40 or fewer endarterectomies per year compared to higher volume hospitals was 2.83 (95% CI = 1.11, 7.19). When volume is treated as a continuous variable in the same model it is also significant (p = .03), and remains so after controlling for teaching status and hospital size.

TABLE 1—Death within 30 Days of Carotid Endarterectomy, Adjusted Odds Ratio and Risk by Patient Subgroup

Patient Age (years)	(N)	Adjusted Odds Ratio* (95% CI)		% Risk (95% CI)
65–69	(733)	1.00		1.1 (0.5, 2.1)
70–74	(671)	2.58 (1.12, 5.95)		2.8 (1.7, 4.4)
75–79	(473)	2.91 (1.22, 6.95)		3.2 (1.8, 5.2)
80+	(212)	4.46 (1.73, 11.50)		4.7 (2.3, 8.5)

Hospital Volume**	(N)	Adjusted Odds Ratio* (95% CI)		% Risk (95% CI)
Over 40	(457)	1.00		1.1 (0.3, 2.5)
21–40	(561)	2.63 (0.94, 7.32)		2.7 (1.5, 4.4)
6–20	(913)	2.90 (1.10, 7.62)		3.0 (2.0, 4.3)
5 or less	(158)	3.14 (0.89, 11.06)		3.2 (1.0, 7.2)
Total	(2089)	—	Average	2.5 (1.9, 3.3)

*Derived from multiple logistic regression with age, sex, surgical volume, and presence or absence of coronary artery disease as covariables. The first level of each category is the reference.

**Volume defined as the number of carotid endarterectomies performed per year on the Medicare population at each hospital.

Risk of Death or Stroke

Because we were unable to characterize the timing in relation to surgery of the majority of strokes recorded on the discharge abstract, we present both a minimum and maximum estimate of post-operative death plus stoke risk in Table 2. Regardless of the definition used, the risk of death or stroke increases significantly with age and at lower volume hospitals.

Individual Hospital Outcomes

We identified all hospitals where the number of observed deaths or definite strokes was greater than what would have been expected ($p < .10$), using as the standard of comparison the outcome rate in those hospitals performing more than 40 carotid endarterectomies per year (Table 3). Twenty-two of the 139 New England hospitals performing endarterectomy were found to have high complication rates compared to high volume hospitals. Only seven New England hospitals performed more than 40 carotid endarterectomies per year on the Medicare population, while 50 hospitals performed five or fewer.

Discussion

Several authors have found no increase in operative mortality in the elderly and have argued that age should not be a consideration when choosing endarterectomy.[21-23] In

TABLE 2—Death or Stroke within 30 Days of Carotid Endarterectomy Adjusted Odds Ratios (AOR)* and Risk by Patient Subgroup

		Death or Definite Stroke		Death or Possible Stroke	
Patient Age (years)	(N)	AOR (95% CI)	% Risk (95% CI)	AOR (95% CI)	% Risk (95% CI)
65–69	(773)	1.00	2.7 (1.7, 4.2)	1.00	8.7 (6.8, 11.0)
70–74	(671)	1.59 (0.89, 2.84)	4.3 (2.9, 6.2)	1.61 (1.15, 2.27)	13.1 (10.7, 15.9)
75–79	(473)	1.87 (1.02, 3.44)	5.1 (3.3, 7.5)	1.54 (1.06, 2.25)	12.7 (9.8, 16.0)
80+	(212)	2.54 (1.25, 5.14)	6.6 (3.7, 10.8)	2.21 (1.42, 3.45)	17.0 (12.2, 22.7)
Hospital Volume**	(N)	AOR (95% CI)	% Risk (95% CI)	AOR (95% CI)	% Risk (95% CI)
over 40	(457)	1.00	2.0 (0.9, 3.7)	1.00	9.2 (6.7, 12.2)
21–40	(561)	2.52 (1.16, 5.44)	4.6 (3.0, 6.7)	1.19 (0.78, 1.80)	10.9 (8.4, 13.7)
6–20	(913)	2.55 (1.23, 5.28)	4.7 (3.4, 6.3)	1.46 (1.00, 2.12)	13.0 (10.9, 15.4)
5 or less	(158)	3.13 (1.22, 8.07)	5.7 (2.6, 10.5)	1.96 (1.15, 3.32)	16.5 (11.0, 23.2)
Total	(2089)	—	4.2 (3.3, 5.1)	—	11.9 (10.5, 13.3)

*Derived from multiple logistic regression with age, sex, surgical volume, and presence or absence of coronary artery disease as covariables. The first level of each category is the reference.

**Volume defined as the number of carotid endarterectomies performed per year on the Medicare population at each hospital.

TABLE 3—Number of Hospitals Identified as Exceeding Outcome Thresholds by Surgical Volume

Volume of Operations per Year	Number of Hospitals in Group	Hospitals Identified with Definite Stroke plus Death Rates Significantly above Expected ($p \le .10$)*
1 to 5	50	6
6 to 20	65	10
21 to 40	17	6
Over 40	7	0
Total	139	22

*For each hospital, the expected proportion of definite strokes plus deaths is assumed to be .0197, the proportion that occurred in the high-volume hospitals. Given this proportion, the probability that a hospital would have at least as many strokes plus deaths as actually reported is computed. A hospital where this probability is .10 or less is considered to have a significantly high rate. The number of definite strokes plus deaths is assumed to have a binomial distribution.

contrast, our study demonstrates a markedly increased mortality risk with increasing age. Two-thirds of the New England Medicare patients undergoing endarterectomy were over the age of 70 and faced an average three-fold increased risk of operative death compared to younger patients. Patients undergoing surgery at low-volume hospitals also faced a three-fold increased risk of death compared to those receiving surgery at high-volume hospitals.

Our analyses of post-operative stroke risk are consistent with the findings of our mortality analyses, but leave uncertainty about the actual risk of non-fatal stroke following endarterectomy. Several lines of evidence suggest that the actual risk is considerably higher than our minimum estimate of death plus stroke risk (4.2 percent). First, a population-based sample of Medicare patients undergoing carotid endarterectomy in 1981 were found to have death plus stroke risks of 9.8 percent overall, based on review of hospital records.[29] Second, each of the five recent series that described the outcomes of carotid endarterectomy across multiple hospitals report a ratio of non-fatal strokes to deaths of approximately two-to-one.[23,29–32] Applying this ratio to our data would suggest a stroke plus death risk of 7.5 percent overall, high enough to raise serious questions about the relative risks and benefits of endarterectomy in the elderly.[29]

A potential criticism of our findings is that the data on disease coded in discharge abstracts may not provide sufficient information to adequately control for severity of illness and comorbidity.[33] Although we share this concern,[27] particularly with respect to the comparison of aggregate hospital mortality rates, the limitations of the diagnostic coding should not undermine our current analysis. It is clear from both the Rand consensus panel's conclusions[12] and from decision analysis[11] that high post-operative complication rates virtually eliminate the potential benefit of the operation for all patient subgroups, no matter what assumptions are made about the operation's efficacy or the stroke rate among medically managed patients. Thus patients who undergo surgery at hospitals which experience high rates of post-operative morbidity and mortality are unlikely to benefit from the surgery, regardless of whether the high rate is due to patient selection or to poor technical performance of the procedure itself. In fact, severity adjustment could obscure important differences in outcomes across institutions due to what may be inappropriate patient selection.

Given the current uncertainty about the value of carotid endarterectomy, the importance of operative risk to the expected benefit of the procedure and the wide variation in post-operative complication rates at individual hospitals, patients and their physicians should seek reliable data on the operative risk of death and stroke at the hospital where the operation is to be performed. These data are currently unavailable. One solution would be for individual hospitals to review their records and report their results. Such a monitoring program would be limited by the degree to which hospitals could ascertain post-discharge events and by concern about the potential conflict of interest for hospitals collecting and reporting the data.

Our study suggests a practical alternative for the 70 percent of patients undergoing endarterectomy who are over the age of 65[1] and thus eligible for Medicare. Data in the claims files could be used to establish an ongoing monitoring program that would identify all patients undergoing endarterectomy and ascertain post-operative mortality. Chart reviews could be carried out to ascertain post-operative stroke rates and accurately classify patients' severity of illness and comorbidity. Accurate risk stratification, whether by age or clinical indications or both, would improve hospital comparisons and enhance individual clinical decision-making. The effort required to establish such a monitoring program would be far less than that required to establish an independent prospective data collection system to provide patients and physicians with accurate information on institution-specific operative risks.

One limitation of any monitoring program is that hospitals where few operations are performed cannot be reliably evaluated; small numbers make it impossible to distinguish the few with excellent outcomes from those with worse than average performance.[34] Given the importance of a low complication rate to the potential benefit of carotid endarterectomy, the difficulty of evaluating low volume hospitals' outcomes and the demonstrated volume-outcome association, it may be reasonable to restrict carotid artery surgery to high volume hospitals that demonstrate excellent outcomes.

The controversy surrounding carotid endarterectomy has been heightened by recent reports confirming that a significant proportion of the carotid endarterectomies performed in the United States are carried out in clinical settings where experts believe the risk outweighs the potential benefit.[11,29] Our findings of increased risk in the elderly and at low-volume hospitals are consistent with this conclusion. An ethical solution to the controversy would be to provide patients with the best available information on the potential benefits and risks of alternative treatments so that the decision may reflect their values.[34] The Medicare claims files, supplemented as necessary with chart review, provide a readily available and feasible method to improve the quality of information available to patients and their physicians.

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Geographic Variation in the Onset of Decline of Male Ischemic Heart Disease Mortality in the Netherlands

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Abstract: We studied variation in the year of onset of ischemic heart disease mortality decline among regions (n = 39) in the Netherlands. Using loglinear regression methods, a quadratic regression model was fitted to the observed numbers of male deaths in each region in the period 1950–84, controlling for changes in age-structure of populations. The quadratic regression model proved inadequate to describe the mortality experience of females.

For the country as a whole, the estimated year of onset of male ischemic heart disease mortality decline is 1973.9. The difference between the earliest and the latest region is almost nine years (1970.0 and 1978.9, respectively). An early onset of decline (≤1972) is only found in the urbanized, western part of the country. A later (≥1975) onset of decline is characteristically found in more peripheral regions in the South-West and South-East, as well as in the North.

Exploratory correlation and regression analyses show that both average income and percent of population living in larger cities have independent, negative associations with the year of onset of male ischemic heart disease mortality decline.

We argue that regional variation in the timing of lifestyle changes is a more plausible explanation of these observations than regional variation in the timing of medical care improvements. (*Am J Public Health* 1989; 79:1621–1627.)

Introduction

After more than 30 years of rising death rates from ischemic heart disease, a dramatic decline has occurred in many industrialized countries, with large differences in the year of onset of decline.[1]

In the United States, the mortality rates plateaued in the early 1960s, and had definitely started to decline in 1968.[2–4] In Australia, death rates from ischemic heart disease reached their highest levels in 1965–67,[5] and in New Zealand in 1968.[6]

In Europe, the decline started in the first half of the 1970s in Denmark, Norway, Finland, the Federal Republic of Germany, Belgium, the Netherlands, and England and Wales. On the other hand, in Ireland, France, Switzerland, Austria, Italy, and Sweden, ischemic heart disease mortality started to decline in the second half of the 1970s. In most Eastern European countries, including Hungary, Poland and Romania, mortality rates were still increasing at that time.[1]

These international differences suggest a process of gradual and worldwide diffusion of the factors responsible for ischemic heart disease mortality decline: from a focus in the New World (including Australia and New Zealand) to Europe, and within Europe from West to East.

Such a process of gradual diffusion has also been found within the United States. Two studies of geographic variation in the year of onset of ischemic heart disease mortality decline within the United States have been reported, showing a considerable range of variation between State Economic Areas[7] and States,[8] respectively. Early declines were reported for California[9,10] and New York.[11] In general, metropolitan areas had earlier years of onset of decline than non-metropolitan areas, suggesting that the factors responsible for the mortality trend reversal have spread across the population according to a general pattern of cultural diffusion.[7]

In the Netherlands, age-standardized mortality from ischemic heart disease reached its highest level in 1972, for both males and females (Figure 1).

Among males, ischemic heart disease death rates more than doubled from 105 per 100,000 person-years in 1950, to 236 per 100,000 person-years in 1972. Thereafter, a gradual

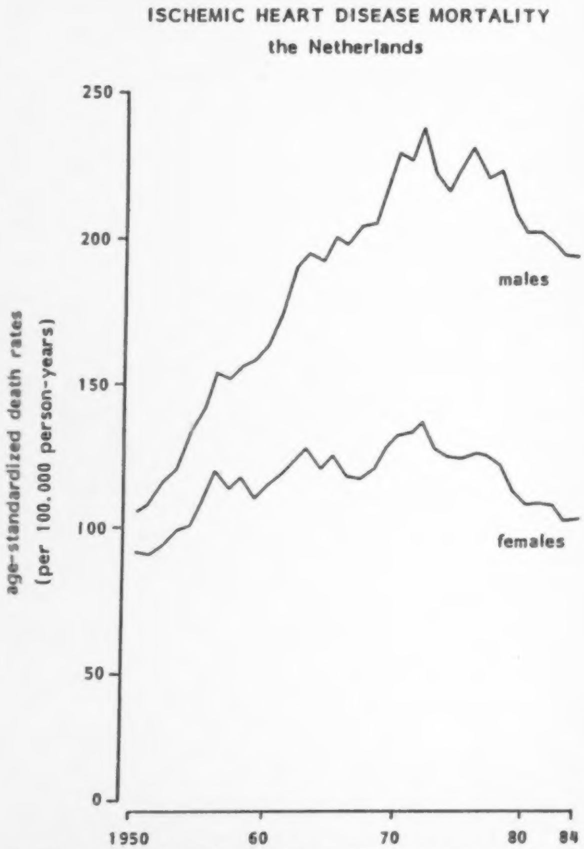

FIGURE 1—Age-Standardized Trends in Mortality from Ischemic Heart Disease, by Sex, the Netherlands, 1950–84

ICD-code numbers	ICD-6/7 1950–68	ICD-8 1969–78	ICD-9 1979–
Ischemic heart disease	420,422.1	410–414	410–414, 429.2

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decline set in, with a secondary peak in 1976, to a rate of 191 per 100,000 person-years in 1984. Among females, the increase between 1950 and 1972 was much less steep than in males, but the decrease after 1972 was almost as rapid.

The causes of this trend reversal have not been extensively investigated in the Netherlands. There is some evidence for a reduction of serum cholesterol between 1960 and 1979.[12,13] The prevalence of self-reported smoking (but not the number of cigarettes sold per capita) has been declining since 1958 in males and since 1970 in females.[14,15] There is no information on trends in the prevalence or the detection and treatment of hypertension in the Netherlands. Changes in the medical management of patients with ischemic heart disease have probably mirrored those in other industrialized countries, but quantitative information on trends at the national level is not available.

The objective of our study of regional variation in the year of onset of mortality decline is to provide some clues to the causes of the mortality trend reversal for ischemic heart disease: if a pattern of geographical variation has been identified, it will be easier to find the factors responsible for the trend reversal, because these must have followed the same pattern. Because of estimation problems in females, we will report only on variation in the year of onset of decline in males.

Methods

Numbers of deaths by age (<1, 1–4, 5–14, ... 85–94, 95+ years), sex, calendar-year (1950–84), region, and four-digit-number of the International Classification of Diseases (ICD) were available from a large computer file supplied by the Netherlands Central Bureau of Statistics.

For statistical purposes, the Netherlands have been divided into 40 regions, each representing a geographical area serviced by one or more larger cities (Figure 2). Total population of the Netherlands was 14.4 million in 1984; the median population size of these 40 regions was approximately 300,000 in 1984.

The economic and cultural center of the country is located in the largely urbanized quadrangle bordered by the regions containing the four largest cities of Amsterdam, The Hague, Rotterdam, and Utrecht. In the rest of the country, a number of secondary centers is found (indicated in Figure 2, which shows the location of other cities with more than 100,000 inhabitants), with other regions occupying a more "peripheral" position, from both a geographic and a cultural point of view. Region no. 40, land which was recently claimed from the waters of the former Zuiderzee, was excluded from the analysis because its population was very small before the end of the 1970s.

For ischemic heart disease, ICD-revisions 6 and 7, we included 420 (arteriosclerotic heart disease) and 422.1 (other myocardial degeneration with arteriosclerosis). The latter number was included because bridge-coding between revisions 7 and 8 revealed that this category was transferred to the ischemic heart disease-section in ICD-revision 8, 410-414.[16] For ICD-revision 9 we included 410-414 (ischemic heart disease) and 429.2 (cardiovascular disease, unspecified), the latter because bridge-coding between revisions 8 and 9 showed that its separation from ischemic heart disease in revision 9 was the main factor behind the lower-than-1 comparability ratio.[17] As Figure 1 shows, the mortality trend shows no signs of sudden changes at the moments of introduction of new ICD-revisions (1969 for ICD-8, and 1979 for ICD-9), suggesting that comparability over time of our cause-of-death category "ischemic heart disease" is satisfactory.

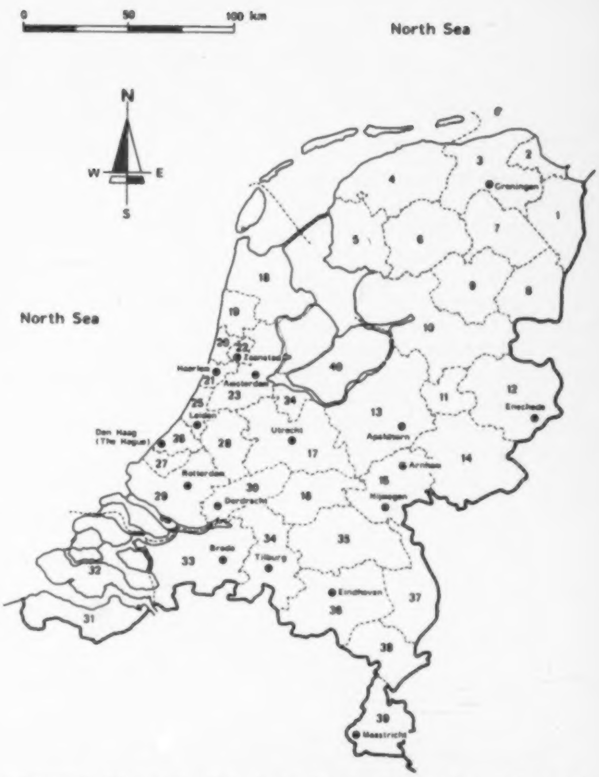

FIGURE 2—Map of the Netherlands, Showing Division into 40 Statistical Regions and Location of Cities with More than 100,000 Inhabitants

Higher age-groups were not excluded from the analysis, because declines in ischemic heart disease mortality have occurred at all ages, and have even been remarkably uniform in the Netherlands.[18]

Population numbers by age, sex, calendar-year, and region were available from several sources based upon the Dutch population register which is kept up to date continuously. Approximate person-years at risk for the deaths occurring in a given calendar-year were calculated by averaging the population numbers for December 31 of both the current and previous years.

Loglinear regression analysis was used to obtain estimates for the years of onset of decline of ischemic heart disease mortality. We chose the loglinear model because our dependent variable is a person-time rate. The model was fit by assuming a Poisson distribution for the number of cases observed. For each region a quadratic regression equation was fit to the observed number of deaths with the following structure:

$$E(y_i) = N_i e^{\alpha_i + \alpha + \beta_1 X + \beta_2 X^2}$$

in which:
$E(y)$ = expected number of deaths
N = person-years at risk
e = base of natural logarithm
α = national mortality level, 1950
α' = regional departure from national mortality level, 1950
β_1, β_2 = regional mortality trend parameters

X = calendar-year
i = subscript denoting age-group.

We used national mortality levels by age (α_i) to adjust the regional trend parameters for changes in age-composition of the population. From β_1 and β_2 the year of maximum mortality rates, marking the onset of decline, can easily be calculated. The variance and covariance estimates of β_1 and β_2 from the regression calculations and a Taylor series expansion[19] were used to obtain variance estimates for the years of maximum mortality. An approximate asymmetric 95 percent confidence interval was derived with the help of Fieller's theorem.[20] The calculations were performed with the GLIM-package.[21]

In order to test the goodness-of-fit of our quadratic regression model, we inspected plots of observed and expected (under the model) numbers of deaths and performed a Wald-Wolfowitz "runs test" on the residuals.[22] Both methods suggested a considerably better fit in males than in females.

In males, the estimated years of onset of mortality decline closely correspond to what one would guess on the basis of visual inspection. At the national level, for example, the estimate is 1973.90, slightly later than 1972 due to the secondary peak in 1976, but still a satisfactory measure. In females, however, the estimated years of onset are considerably earlier than 1972, due to the flat rise and steep decline of the curve (Figure 1). At the national level the estimate is 1969.03.

The "runs test" identified only two regions with statistically significant ($p < .05$) deviations from a quadratic model in males (to be expected from chance alone), but four in females. In these cases, fitting higher order polynomials produced more satisfactory fits, but did not change the estimated year of onset of decline in males. In females, however, it produced very different estimates, suggesting that the quadratic regression model is really inadequate. For these reasons we decided to restrict this presentation to the results obtained for males.

The national number of deaths from ischemic heart disease among males over the entire period 1950–84 was 412,613. Regional numbers varied from 1,744 to 41,785, with 8,121 deaths in the region with the median population size.

Results

The estimated years of maximum male mortality from ischemic heart disease, marking the year of onset of mortality decline, are presented in the appendix, together with their 95% confidence intervals. Even in a small country like the Netherlands, the difference between the earliest and the latest region is almost nine years: the range of variation is from 1970.0 to 1978.9. Although confidence intervals are sometimes wide, especially for smaller regions, there is no overlap between the confidence intervals for regions with an early onset of decline ($\leq$1972) and those with a late onset of decline ($\geq$1977).

There appears to be a bimodal frequency distribution of the year of onset of decline of ischemic heart disease mortality among males in Dutch regions (Figure 3). The first peak in the distribution represents 17 regions (containing 50 percent of the Dutch male population) with an estimated year of onset of decline in 1973 or 1974, i.e., corresponding to the national average. Only six regions (containing 26 percent of the Dutch male population) had years of onset of decline before 1973. After 1974, a smaller second peak in the frequency distribution appears in 1977. This second peak may partly explain the second peak in the national mortality trend in 1976 (Figure 1), although the years do not exactly correspond.

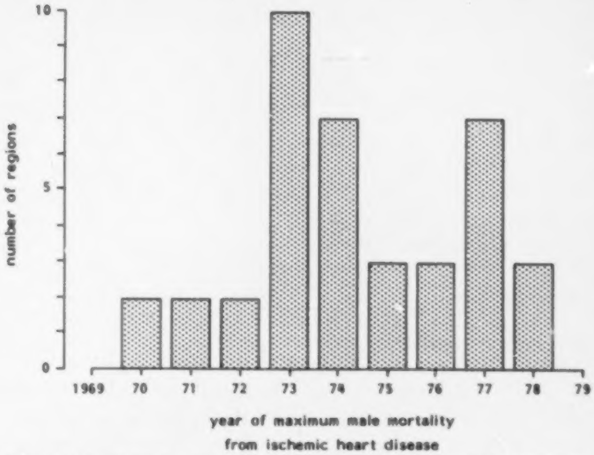

FIGURE 3—Frequency Distribution of the Year of Maximum Male Mortality from Ischemic Heart Disease, 39 Regions, the Netherlands

Figure 4 illustrates the spatial distribution of the year of

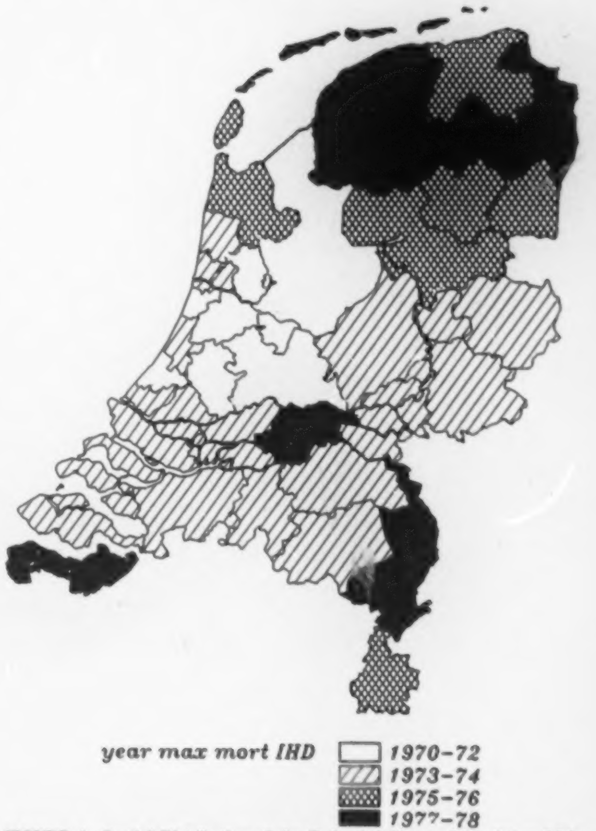

FIGURE 4—Spatial Distribution of the Estimated Years of Maximum Male Mortality from Ischemic Heart Disease, the Netherlands

onset of decline of ischemic heart disease mortality among males in the Netherlands.

An early onset of decline (≤1972) is only found in a number of regions located in the urbanized, western part of the country, including the regions containing Amsterdam and the Hague, but not the region containing Rotterdam. A later (≥1975) or late (≥1977) onset of decline are characteristically found in more peripheral regions in the South-West and South-East, as well as in the North. There is only one exception: region no. 16, very centrally located from a purely geographical point of view but rural in character, had its year of onset of mortality decline in 1977.

Regional differences in the year of onset of mortality decline represent important differences in the timing of the ischemic heart disease mortality "epidemic" as a whole. Figure 5 shows two "cross-sectional" views of ischemic heart disease mortality among males, for the early 1950s and early 1980s, respectively. In 1980–84, the regions in the highest quintile of the distribution were mainly located in the North and South-East, parts of the country which were also characterized by a late onset of mortality decline.

This pattern gradually emerged from a completely different pattern of variation in the early 1950s, when areas with higher than average mortality were mainly found in the West. (The product-moment correlations between year of onset of decline and mortality levels in 1980–84 and 1950–54 were 0.69 and −0.48, respectively.) This suggests that the epidemic of ischemic heart disease mortality started and ended earlier in the western parts of the Netherlands.

In an exploratory search for covariates of regional differences in the year of onset of male ischemic heart disease mortality decline we calculated product-moment correlation coefficients with a number of sociodemographic variables, all measured in a year close to (mostly just before) 1973 (Table 1).

The correlations confirm the impression that lower values of year of onset of decline are found in more urbanized areas: the correlation coefficient with percentage population living in larger cities is 0.74. On the other hand, the year of onset of decline also appears to be strongly associated with socioeconomic characteristics: lower average income goes together with a later onset of mortality decline.

Two variables measuring another important cultural division, percentage population of protestant c.q. roman-catholic religious affiliation, usually powerful predictors of mortality patterns in the Netherlands, are not associated with the year of onset of ischemic heart disease mortality decline.

The results of a multiple regression analysis (Table 2) show that both average income and percentage living in larger cities have an independent, negative association with year of onset of decline. After controlling for these two powerful predictors two additional socioeconomic variables appear to be associated with

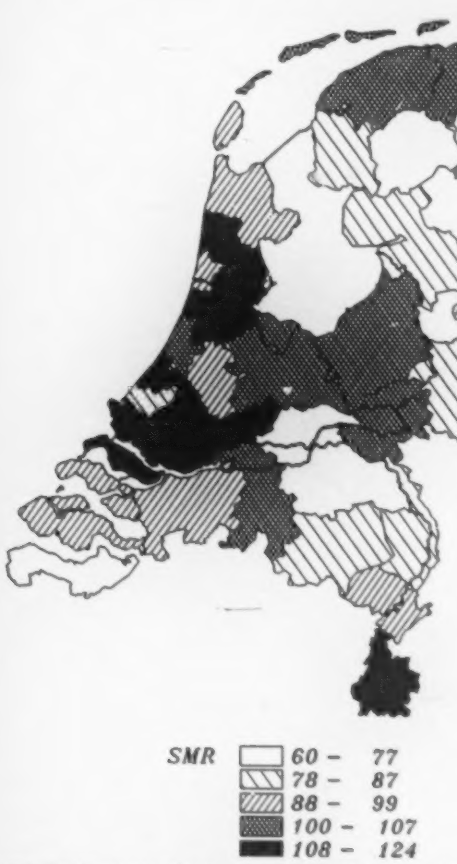

FIGURE 5a—Spatial Distribution of Age-Standardized Ratios of Male Ischemic Heart Disease Mortality, the Netherlands, 1950–54

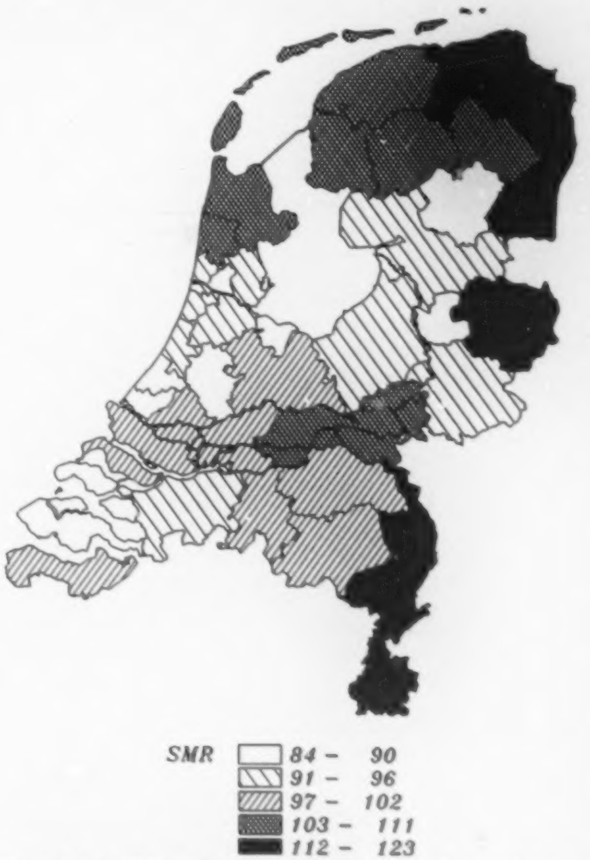

FIGURE 5b—Spatial Distribution of Age-Standardized Ratios of Male Ischemic Heart Disease Mortality, the Netherlands, 1980–84

TABLE 1—Correlations between the Year of Onset of Male Ischemic Heart Disease Mortality Decline and Sociodemographic Variables, the Netherlands

Variables[a]	Mean	Coefficient of Variation	Simple Correlation Coefficient[b]
Average income per inhabitant (guilders, in 000s)	4.95	0.13	−0.79
% Working population unemployed	1.14	0.40	0.32
% Economically active population working in agriculture	2.09	0.80	0.34
% Economically active population working in industry	38.4	0.26	0.23
% Population living in larger cities	23.5	1.00	−0.74
% Population of protestant denomination	37.5	0.55	0.25
% Population of roman-catholic denomination	37.0	0.77	−0.08
% Population without Dutch nationality	1.73	0.55	−0.26

[a] measured 1969–72

[b] product-moment correlation coefficient r

TABLE 2—Regression of the Year of Onset of Male Ischemic Heart Disease Mortality Decline on Sociodemographic Variables, the Netherlands

Variable[a]	Multiple Regression Coefficient[b]	95% Confidence Interval
Average income	−1.693	−2.695, −0.691
% Living in larger cities	−0.046	−0.070, −0.022
% Working in industry	−0.053	−0.096, −0.010
% Unemployed	1.122	0.168, 2.076

$R^2 = 0.773$

[a] Variables were selected from those included in Table 1 in a forward stepping procedure on the basis of F-tests, and are mentioned in the order of inclusion into the regression equation.

[b] Ordinary least squares regression coefficient β.

the year of onset of decline: percentage working in industry (negative association) and percentage unemployed (positive association). The percentage of variance explained by these four variables is considerable, 77 percent.

Discussion

In their analysis of the onset of decline of ischemic heart disease mortality in the United States, Ragland and associates found a range of variation between States of almost 15 years for White males aged 55–64 years, from 1952.82 in California to 1967.29 in West Virginia.[8] If their units of analysis had been smaller, they would probably have found an even wider range of variation. This is evident from the study of State Economic Areas by Wing and associates, who found that 1.6 percent of these areas had a year of onset of decline as late as in or after 1973.[7] We feel that a range of variation of almost nine years between regions in such a small country as the Netherlands is no less dramatic.

Although differences in cause-of-death certification are a potential threat to the validity of all studies of geographical variation in cause-specific mortality, it is implausible that such differences could account for the variation in year of onset of decline of ischemic heart disease mortality observed in this study. As changes in certification have probably not contributed much to the trend reversal of ischemic heart disease mortality at the national level,[4] it is difficult to see how any regional differences in such changes could have produced such dramatic variation in year of onset of decline. Moreover, the validity of Dutch cause-of-death certification compares favorably with that of many other European countries.[23,24] Coding of death certificates is done centrally in the Netherlands.

Another potential threat to the validity of our study might arise from violations of the assumptions underlying the quadratic regression model (i.e., a unimodal distribution of mortality over time, and equal rates of increase before and decrease after the top-year). These assumptions were reasonably satisfied for males, however. Although the national mortality trend for males is bimodal (Figure 1), this appears to be the result, at least in part, of the superposition of two unimodal trends with a different timing (Figure 2). For females, the rate of increase before the early 1970s is much smaller than the rate of decrease thereafter, so that a quadratic regression model cannot be used to estimate years of maximum mortality. This is unfortunate, because a comparison of regional patterns for females with those for males could give important additional information. Further study of female ischemic heart disease mortality decline in the Netherlands, using an appropriate model form, is therefore to be recommended.

Despite a vast amount of literature, there is still no consensus on the relative contributions of changes in risk factor levels and improvements in medical care to the reduction of ischemic heart disease mortality.

A decline in incidence of myocardial infarction has been found in a number of population-based studies, both from the United States[25–27] and from a few other countries.[28–30] Lifestyle changes, leading to reduced tobacco consumption and lower serum cholesterol levels, have almost certainly contributed to this decline.[4,6,31,32]

Examinations of trends in myocardial infarction survival have produced a more diffuse picture than those of trends in incidence. Some studies have reported a decrease in in-hospital case fatality rates[26,27] but others have not done so.[33,34] Studies which have examined long-term survival reported no improvements.[26,35,36] (In most of these studies, the time-periods examined preceded widespread use of beta-blocking agents for long-term patient management.) Nevertheless, a recent comparison of two cohorts suggested that early detection and treatment did contribute to the observed decline in ischemic heart disease mortality.[37] Estimates of the contribution of specific medical interventions to the decline in the United States between 1968 and 1976 range as high as 40 percent.[32]

Although the pattern of international variation in onset of decline mentioned in the introduction could be consistent with a diffusion process of either lifestyle or medical care innovations, a detailed comparison of available data suggests that lifestyle factors are the better candidates.[38] Sweden is a particularly enlightening case because it combines a very advanced health care system with a late year of onset of decline of ischemic heart disease mortality.[39]

It is important to note that the spatial distribution of the year of onset of ischemic heart disease mortality decline within the Netherlands is analogous to that observed in the

United States.[7,8] In both countries the decline started in the highly urbanized areas which are regarded as the cultural center(s) of the nation, and gradually diffused to more "peripheral" regions. In the Netherlands, as in the United States, the factors responsible for the mortality trend reversal thus seem to have spread according to a general pattern of cultural diffusion. In the Netherlands, as in the United States,[40] socioeconomic factors seem to play an important role in this diffusion process. Theoretically, both changes in medical care (improved control of hypertension and/or better management of patients with cardiovascular disease) and changes in lifestyle could have spread according to such a pattern. For a number of reasons, however, we believe that differences in timing of lifestyle changes are a more plausible explanation for the Netherlands:

- If differences in timing of changes in medical care were to explain the regional variation in onset of decline of ischemic heart disease mortality, one would have to assume that the providers of medical care in the most "backward" regions adopted these improvements almost nine years later than those in the most "advanced" regions. In view of the short lines of communication within the Dutch medical care system, and of the geographically homogeneous insurance schemes, this seems a very unrealistic assumption. Diffusion of changes in lifestyle in the population at large, on the other hand, is probably a much slower process.
- If the degree of urbanization were an indicator of early diffusion of innovations in the medical care of ischemic heart disease, one would also expect an association between the presence of a university hospital, as a typically metropolitan and diffusion-prone institution, and the year of onset of mortality decline. In 1970, there were seven university hospitals in the Netherlands, all located in cities with 100,000+ population. The average year of onset of male ischemic heart disease mortality decline in regions with a university hospital was 1973.6, almost exactly the same as the national average. This is due to the fact that one university hospital is located in the North (region no. 3, with its year of onset of decline in 1976), and only three are located in regions with an early decline (one in no. 17 and two in no. 23, with their years of onset of decline in 1972 and 1971, respectively). Also after controlling for average income, there is no association between the percentage of hospital beds located in university hospitals and the year of onset of decline.
- The association between year of onset of decline and average income suggests that an early adoption of factors promoting ischemic heart disease mortality decline is not only determined by circumstances directly linked to living in large cities, but also to the more favorable socioeconomic circumstances prevailing in the western parts of the country. This suggestion is supported by the results of analyses at the individual level from England and Wales, which have shown that ischemic heart disease mortality decline started earlier in higher socioeconomic groups.[41] In England and Wales, this variation corresponds to differing trends in lifestyle related risk factors such as smoking and high cholesterol intake.[42] In the Netherlands, too, the decline in smoking among males started in the higher socioeconomic groups.[14]

APPENDIX

Estimated Years of Maximum Male Mortality from Ischemic Heart Disease, by Region, the Netherlands

Region[a]	Top-year	95% confidence intervals
1	1977.2	1975.8, 1979.0
2	1977.5	1974.4, 1984.5
3	1976.9	1975.6, 1978.7
4	1977.1	1974.7, 1978.0
5	1978.0	1975.5, 1982.4
6	1977.8	1976.1, 1980.5
7	1977.7	1975.8, 1980.6
8	1976.2	1974.9, 1977.9
9	1976.5	1974.7, 1979.4
10	1975.1	1974.0, 1976.6
11	1974.2	1973.0, 1976.0
12	1973.7	1973.2, 1974.4
13	1974.9	1973.8, 1976.4
14	1974.8	1973.7, 1976.3
15	1973.6	1972.9, 1974.3
16	1977.0	1975.3, 1979.6
17	1972.9	1972.3, 1973.7
18	1975.7	1974.3, 1977.6
19	1974.8	1972.8, 1987.2
20	1973.0	1971.7, 1974.8
21	1970.2	1969.4, 1971.2
22	1973.2	1971.2, 1976.5
23	1971.5	1971.1, 1972.0
24	1971.4	1970.5, 1972.5
25	1973.6	1972.3, 1975.3
26	1970.0	1969.6, 1970.5
27	1973.6	1972.4, 1975.3
28	1972.7	1971.7, 1973.9
29	1973.2	1972.7, 1973.9
30	1974.0	1973.0, 1975.4
31	1977.2	1974.8, 1981.3
32	1974.3	1972.8, 1976.5
33	1973.7	1973.0, 1974.6
34	1973.1	1972.3, 1974.0
35	1974.2	1973.5, 1975.0
36	1973.4	1972.8, 1974.0
37	1978.3	1976.5, 1981.1
38	1978.9	1976.7, 1982.6
39	1975.0	1974.3, 1975.9
Netherlands	1973.9	1973.7, 1974.1

[a] see Figure 2

ACKNOWLEDGMENTS

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Effect of a Gatekeeper Plan on Health Services Use and Charges: A Randomized Trial

DIANE P. MARTIN, PHD, PAULA DIEHR, PHD, KURT F. PRICE, MS, AND WILLIAM C. RICHARDSON, PHD

Abstract: A randomized trial was conducted to determine the effectiveness of a health care plan which uses physicians as gatekeepers to control health services use and charges. New enrollees in United Healthcare (UHC), an independent practice association, were randomly assigned to the standard UHC plan requiring a gatekeeper, or to an alternate plan with equal benefits but without a gatekeeper. Individuals in both plans were similar in demographic characteristics, perceived health status, and other health insurance coverage. The gatekeeper plan had 6 percent lower total charges per enrollee than the plan without a gatekeeper. There were minor differences in hospital use and charges. Ambulatory charges were $21 lower per person per year in the plan with a gatekeeper (95% CI = −39.9, −2.1) and these were due to .3 fewer visits to specialists (95% CI = −0.50, −0.10). We conclude that a health plan which incorporates incentives and penalties for physicians to act as gatekeepers can reduce the cost of ambulatory services by limiting specialist visits. (*Am J Public Health* 1989; 79:1628–1632.)

Introduction

Many health plans use a variety of mechanisms to encourage primary care physicians to act as "gatekeepers" in managing the medical care of their patients.[1,2] Incentives and penalties presumably will control cost by changing the gatekeeper's own practices and by limiting inappropriate hospital and referral care. The effectiveness of gatekeeper plans depends on the ability of participating physicians to coordinate the medical care system efficiently.

Gatekeepers are used by many health maintenance organizations (HMOs), independent practice associations, preferred provider organizations, and some Medicaid plans,[3,4] but evidence that they save money is lacking.[5] Descriptive data exist showing that gatekeepers can decrease the use of services if the physicians share financial risk of providing care,[6-8] but there are no data from controlled studies.

We therefore conducted a randomized trial of a plan using primary care physicians as gatekeepers to control health services use and charges. We randomly assigned study participants to United Healthcare, an independent practice association using gatekeepers, or to Northwest Healthcare, a plan with the same benefits but no gatekeeper. Because the economic incentives in the gatekeeper plan encouraged use of the gatekeeper and discouraged hospital and referral care, we expected that United Healthcare would have fewer hospitalizations, fewer visits to specialists, and lower use of procedures.

Methods

Population and Random Assignment

United Healthcare (UHC) was one of three health care plans available to Washington State employees in the Seattle metropolitan area during the 1979 open enrollment. All subscribers (n = 1,239) and their family members who enrolled in UHC for the first time were asked to participate in the experiment and complete an initial telephone interview. Retirees were excluded from the study.

The 90 percent of subscribers (families) who agreed to participate were assigned to a health plan according to a randomization scheme stratified by previous health care plan and type of employer, because State health insurance information showed that use of services differed significantly for these subgroups.[9] Previous health care plans were: Blue Cross of Washington and Alaska, which provided comprehensive inpatient and outpatient coverage; Group Health Cooperative of Puget Sound, a staff model HMO; or no previous State health plan, for new employees. The two types of employer were: an institution of higher education, or other State agency. We assigned 555 subscribers and their dependents (1,419 enrollees) to the standard United Healthcare plan with a gatekeeper and 558 subscribers (1,408 enrollees) to the plan without a gatekeeper, Northwest Healthcare.

The Two Plans

Both plans were offered by the SAFECO Insurance Company and had the same insurance premiums and benefits. Premiums for the subscriber and all dependents were paid in full by the State. Both plans provided inpatient, outpatient, and preventive care coverage with copayments and limits on maternity and mental health services. There was no deductible or coinsurance, but there were copayments of $2 for each prescription and $25 for each general physical examination.

The plan with a gatekeeper, United Healthcare, attempted to contain costs by enlisting primary care physicians in private practice as gatekeepers. The details of this plan have been published earlier.[10] When the present study began, most gatekeepers had been in the plan for two years. Hospitalization and other major procedures entailed no utilization review, prior authorization, or mandated second opinions, and there was little emphasis on quality assurance or provider education. Because gatekeepers had only 25 plan patients on average, capitation reimbursement was not practical, and almost all were reimbursed at 95 percent of their charges. Thus, primary care physicians had no financial incentive to limit costs of the care they provided.

There were stronger incentives for gatekeepers to monitor referral care. In 1979, referral care consumed 74 percent of the amount paid for health services by United Healthcare.[9] A risk-sharing account was established for each gatekeeper based on the age and sex of his or her panel of enrollees. These funds were then used to cover hospitalization, drugs, specialty, ancillary, and other referral costs. Catastrophic

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costs greater than $5,000 per patient were paid by a reinsurance program. At the end of the year, the gatekeeper shared with United Healthcare any deficit or surplus remaining in the account. If there was a surplus in the account at the end of the year, it was split equally between the gatekeeper and UHC. If there was a deficit, the gatekeeper had to pay back what was over-spent, up to an amount equal to 10 percent of his fee-for-service charges. The financial risk was not great to the physician. If the account was overdrawn, the minimum reimbursement rate was 85 percent (95 percent initial payment minus 10 percent), which was considered to be a reasonable return on charges. Furthermore, the risk applied to only a small proportion of the physician's total practice.

UHC gatekeepers could refer patients to the specialists or hospitals of their choice. Referrals could be made only to a few types of non-physician health professionals such as optometrists, psychologists, chiropractors, and occupational and physical therapists. There were no financial incentives for specialists, non-physicians, or hospitals to practice efficient care, because they were paid 100 percent of their fee-for-service charges.

United Healthcare required each enrollee to select a primary care physician from a list (N = 391) of family physicians (46 percent), internists (22 percent), pediatricians (19 percent), obstetrician/gynecologists (11 percent), and the few general surgeons (2 percent) who had joined the plan as gatekeepers. The gatekeeper was responsible for providing primary care and coordinating referrals for specialty care, ancillary services, and hospitalization. If the enrollee received care from a specialist or visited an emergency room for non-urgent care without approval from the gatekeeper, UHC would not pay the claim and the enrollee was responsible for payment.

Enrollees in the plan without a gatekeeper, Northwest Healthcare, were not required to choose a gatekeeper or to obtain approval for hospitalization, emergency, or specialist visits. These enrollees could receive care from any physician or non-physician health care provider in the Seattle area and the plan would reimburse the provider 100 percent of charges. Many of the same physicians who joined United Healthcare as risk-sharing gatekeepers also saw patients in Northwest Healthcare and were reimbursed in full for charges by that plan. Whereas physicians' offices submitted all claims to United Healthcare, Northwest Healthcare enrollees were required to submit claims to the plan for payment. Northwest Healthcare enrollees were provided with addressed, stamped envelopes to facilitate claims submission.

Data Collection

The adult who made most of the health care decisions for the family was asked to complete a baseline telephone interview. Information obtained from the interview included family composition, demographic information, perceived health status, health insurance coverage, and usual source of health care. This information was used to stratify the random assignment to plans and assess the similarity of the two groups. Enrollees in both plans were followed for 12 months. Data on the use of health services and charges were collected from SAFECO computerized membership and claims records. At the end of the study, a questionnaire was mailed to participants to assess experience with the plans and other health insurance during the previous year and to ascertain out-of-pocket expenditures and satisfaction with the plans. The response rate for the mail questionnaire was 95 percent for each plan.

Charges and Use of Services

We examined charges, percentage of enrollees using any services, and volume of services overall and separated according to primary care physicians versus specialists; ambulatory versus hospital care; type of procedure; and diagnostic category. We partitioned the enrollees into non-users and users of services as described by Duan, *et al*,[11] to separate differences in access to care from the effect of gatekeeper control once someone had entered the system.

Total billed charges for 12 months were calculated for each enrollee, including all charges by primary care providers, specialists, hospitals, and ancillary services. Billed charges, rather than the amount paid by SAFECO, were used to ensure comparability between plans. Payments for coordination of benefits were not taken into account.

We measured the percentage of enrollees using any services or a specific service during the study. We also measured the volume of services used per enrollee or per user, e.g., number of ambulatory care visits, hospital admissions and days.

Definition of Primary Care Physician

A physician was designated as a primary care physician (PCP) if he or she was a general or family practitioner, osteopath, pediatrician, general internist, or obstetrician/gynecologist. All other physicians were defined as specialists. Although we tested two other PCP definitions with similar results, we used this definition to make primary care physicians as similar as possible in the two plans. This was the most conservative definition because it tended to overestimate primary care physicians in the plan without a gatekeeper and therefore any differences in primary care physician use between the two plans would be underestimated.

Characteristics of Enrollees

Data obtained during the baseline telephone interview showed enrollees in the two plans to be similar in demographic characteristics, social structural variables, other insurance coverage, perceived health status, and usual source of health care (Table 1). Enrollees in the gatekeeper plan were more likely to have had internists, obstetrician/gynecologists, or other specialists as their usual source of care. Assuming that care by internists and specialists is more costly than care by family physicians,[12] this difference between groups might reduce the difference in cost between the two plans. On the other hand, internists or obstetrician/gynecologists might be able to handle problems that a family physician would refer to a specialist.

Analysis

The differences in the means or proportions of outcome measures and the 95% confidence intervals around those differences were calculated. Analysis of covariance was used to adjust for the control variables. In general, the plan differences remained significant after controlling for main and independent variables for all enrollees and for users.[9] Although there were few differences between the plans on key independent variables, many independent variables were significantly related to charges in both plans. Including these independent variables in the analysis did not change the findings of the univariate analysis in any substantial way, confirming the success of the randomized design. Because the adjusted and unadjusted results were essentially the same, the unadjusted rates are presented in this paper.*

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*Data available on request to author.

TABLE 1—Selected Characteristics of Enrollees in Gatekeeper or No Gatekeeper Plans*

Characteristics	Gatekeeper (n = 1419)	No Gatekeeper (n = 1408)
Age (years)		
0–5	9.2%	8.3%
6–17	24.3	24.6
18–44	52.6	50.7
45+	13.7	16.4
Mean (years)	26.9	27.7
S.D.	15.8	16.0
Percent Female	53.8%	53.7%
Relationship to Subscriber		
Subscriber	39.1%	39.6%
Spouse	22.6	21.3
Dependent	38.3	39.1
Family Size		
Mean number	2.6	2.5
S.D.	1.5	1.5
Subscriber Years of Education		
Mean years	15.6	15.7
S.D.	3.2	3.1
Spouse Years of Education		
Mean years	15.1	14.8
S.D.	2.9	2.9
Family Income		
Under $10,000	15.1%	16.2%
$10,000–20,000	43.1	41.4
$20,000–30,000	25.8	26.1
Over $30,000	16.0	16.4
Other Health Insurance Coverage	27.9%	27.5%
Perceived Health Status		
Excellent	73.6%	74.0%
Good	23.0	23.4
Fair	3.3	2.2
Poor	.1	.4
Had a Usual Source of Care before Study Period	82.9%	84.2%
Usual Source of Care Was a UHC gatekeeper	45.4%	49.5%
Specialty of Usual Source of Care before Study Period		
Family/General Practitioner	49.6%	57.1%
Pediatrician	22.7	21.7
Internist	13.9	9.5
Obstetrician/Gynecologist	8.0	6.2
Other	5.8	5.5
Months Membership in Plan During Study		
Mean months	10.5	10.3
S.D.	2.2	2.3

*Other characteristics which were measured but are not presented here due to small differences between plans include: marital status, family composition, family health risk scale, employment status, months residence in community, and type of health insurance coverage.

Results

Use and Charges for All Enrollees

Access to care, as reflected by the percentage of enrollees using any services or having at least one ambulatory visit, was high in both plans (Table 2). People in the gatekeeper plan had more visits to a primary care physician (difference = 6%, 95% CI = 2.43, 9.57) and fewer visits to a specialist (difference = −9%, 95% CI = −12.28, −5.72). Ambulatory charges per enrollee were $21 lower in the gatekeeper plan (95% CI = −39.9, −2.1, Table 3), due to significantly fewer visits to specialists.

There were minor differences between plans in hospital use and charges per enrollee (Tables 2 and 3). The gatekeeper plan had lower hospital use than the no gatekeeper plan (Table 4). The plans had similar inpatient charges for specific services, such as medical, surgical, laboratory, and radiological procedures (data not shown).*

Charges and Visits for Ambulatory Care

We examined the effect of the gatekeeper plan among those people who used ambulatory services. PCP ambulatory visits and charges for users were similar, but users in the gatekeeper plan had significantly fewer visits to specialists and lower specialist charges (Table 5).

There were no differences between plans in primary care physicians' use of surgical, laboratory, or radiological procedures, as defined by categories of Current Procedural Terminology codes (data not shown). PCPs in the gatekeeper plan ordered or performed fewer medical procedures on their patients than did PCPs in the no gatekeeper plan (difference = −4%, 95% CI = −5.13, −2.87). Specialists in both plans had similar use and charges for various types of procedures.

We also compared ambulatory use between plans for 20 common diagnostic categories and found that people in the gatekeeper plan had fewer visits for mental health problems, back problems, and acute and chronic respiratory problems, and more of their visits for these problems were with a PCP rather than a specialist (data not presented).*

Experience Using the Plans

The end-of-study survey showed no difference between plans in enrollees' use of other health insurance plans, in the proportion of claims submitted, or in out-of-pocket expenditures during the study period. Enrollees in the gatekeeper plan more often identified internists, obstetrician/gynecologists, and specialists as their usual providers of care, while enrollees in the no gatekeeper plan had more often named family physicians (the distribution is similar to source of care before the study, Table 1). Enrollees were more satisfied with the no gatekeeper plan than the gatekeeper plan (very satisfied: 75% − 64% = 11%, 95% CI = 7.63, 14.37).

TABLE 2—Percentage of Enrollees Using Services in the Gatekeeper and No Gatekeeper Plans

Percentage of Enrollees With:	Gatekeeper (n = 1419)	No Gatekeeper (n = 1408)	Difference (95% CI)
	%	%	
Any Services	74	77	−3 (−6.17, +0.17)
≥1 Ambulatory Visit	68	70	−2 (−5.41, +1.41)
≥1 Primary Care Physician Visit	65	59	+6 (+2.43, +9.57)
≥1 Specialist Visit	23	32	−9 (−12.28, −5.72)
≥1 Hospitalization	4.5	5.0	−0.5 (−2.07, +1.07)

TABLE 3—Annual Charges and Visits for Enrollees in the Gatekeeper and No Gatekeeper Plans

	Gatekeeper (n = 1419) Mean	No Gatekeeper (n = 1408) Mean	Difference (95% CI)
Total Charges	$239	$254	−15 (−73.0, +43.0)
Primary Care Physician	85	96	−11 (−24.4, +2.4)
Specialist	155	158	−3 (−52.8, +46.8)
Hospital Charges	$ 93	$ 87	+6 (−44.9, +56.9)
Primary Care Physician	9	16	−7 (−15.4, +1.4)
Specialist	84	71	+13 (−32.2, +58.2)
Ambulatory Charges	$146	$167	−21 (−39.9, −2.1)
Primary Care Physician	76	80	−4 (−13.2, +5.2)
Specialist	70	87	−17 (−30.1, −3.9)
Ambulatory Visits	2.7	3.0	−0.3 (−0.60, +0.02)
Primary Care Physician	2.1	2.0	+0.1 (−0.12, +0.32)
Specialist	.63	.93	−0.3 (−0.50, −0.10)

TABLE 4—Hospital Use for Enrollees in the Gatekeeper and No Gatekeeper Plans*

	Gatekeeper Mean	No Gatekeeper Mean	Difference (95% CI)
Hospital Admissions per 1000 Person Years	49	56	−7 (−25.1, +11.1)
Hospital Days per 1000 Person Years	205	277	−72 (−207.8, +63.8)
Average Length of Hospital Stay	4.2	4.9	−0.7 (−2.7, +1.3)
Hospital Charges			
Primary Care Physician	$ 196	$ 320	−124 (−280.4, +32.4)
Specialist	$1867	$1432	+435 (−378.9, +1248.9)

*The gatekeeper plan had 64 persons with a total of 70 hospitalizations, and the no gatekeeper plan had 70 persons with a total of 79 hospitalizations.

TABLE 5—Charges and Visits for Ambulatory Users in the Gatekeeper and No Gatekeeper Plans

Per User Per Year	Gatekeeper (n = 1052) Mean	No Gatekeeper (n = 1087) Mean	Difference (95% CI)
Ambulatory Charges	$197	$216	−19 (−42.4, +4.4)
Primary Care Physician	102	104	−2 (−13.3, +9.3)
Specialist	95	113	−18 (−34.8, −1.2)
Ambulatory Visits	4.0	4.3	−0.3 (−0.72, +0.12)
Primary Care Physician	3.1	2.9	+0.2 (−0.08, +0.48)
Specialist	.9	1.3	−0.4 (−0.67, −0.09)

Although satisfaction differed, similar percentages of families switched to other plans during the next open enrollment.

Discussion

In this randomized trial, the gatekeeper plan had lower ambulatory charges per enrollee, primarily due to lower use of specialists. The gatekeeper plan had little impact on hospital use or charges perhaps because of small numbers. The total charges per enrollee were 6 percent lower in the plan with a gatekeeper, a modest difference that was not statistically significant.

The gatekeeper plan required patients to see a primary care physician first, so it was not surprising to find that more enrollees in the gatekeeper plan had a visit to a primary care physician. The gatekeeper plan provided no incentives for PCPs to reduce visits to themselves, and indeed PCP visits and charges were similar in the two plans. There were incentives for gatekeepers to limit referrals and this apparently resulted in lower expenditures for specialist services in the gatekeeper plan. Specialists were offered no incentives by the gatekeeper plan to reduce costs, and once patients were referred to specialists, their volume and charges for various types of procedures were similar in both plans.

A gatekeeper effect was seen for certain diagnoses for which treatment is at the discretion of the patient and physician, such as mental health problems and back problems. We do not know if these differences result from: self-selection, decreased initial visits by the enrollee, differences in diagnostic coding, or the practice patterns or specialty of the regular doctor. Future studies of gatekeeper plans should explore utilization by diagnostic category in more detail.

True gatekeeper effects are probably at least as large as those reported here. First, the population consisted of subscribers and their family members who chose United Healthcare rather than one of the other state health insurance options available to them. Because the pre-enrollment information described the standard UHC plan, all study participants may have been oriented to use PCPs more than the general population; therefore, the differences between the plans may be underestimated. Second, differences between plans may have been underestimated because people in the no gatekeeper plan were more likely to have a regular doctor who was already participating in United Healthcare as a gatekeeper.

Two other issues should be considered when interpreting these results. First, we could not examine the gatekeeper plan effect in more detail because we were using claims data. We have no indication whether a specific visit was initiated by the patient or provider, or whether the patient was seeing a particular provider as a primary care physician or as a specialist. Furthermore, claims data did not allow us to describe how or to what extent a primary care physician was functioning as a gatekeeper. Second, this 12-month study period may not be representative of the plans' or enrollees' experience over time. One possibility is that individuals in the no gatekeeper plan may have used services differently because they knew that the study period was limited to one year. Another possibility is that the "learning period" for new enrollees may have affected the results. We would expect this learning period to be more important for the gatekeeper plan because enrollees had to adjust to more restrictions on their use of services than in the no gatekeeper plan.

Other studies have reported mixed results of the gatekeeper effect on the use of services. Catlin, *et al*, surveyed use of primary care physicians in staff and group model HMOs and excluded independent practice associations (IPAs).[6] They found no significant differences in hospital or ambulatory use among HMOs using more primary care physicians, but the financial incentives offered to primary care physicians were not described. Burkett's descriptive study of an IPA-HMO with gatekeepers and financial incentives similar to UHC found the rate of referral visits to be low.[8] In comparing his results to ours, primary care visits per person per year appeared to be slightly higher and use of referral visits even lower than those in this study.

The findings of another study of UHC conducted at about the same time provide additional evidence for assessing the gatekeeper concept. In that study, we compared the standard UHC gatekeeper plan with two other health plans: Group Health Cooperative, a staff model HMO with primary care case management; and Blue Cross, which provided fee-for-service hospital and ambulatory care. Use and estimated expenditures were compared across the three plans for a different cohort of Washington State employees which was followed for two years.[9] Initial access to care was high at both UHC and Group Health. UHC had lower expenditures for specialists than the other plans, but the savings were not sufficient to decrease overall expenditures. Total annual expenditures per person were similar for UHC and Blue Cross and were significantly higher than expenditures for Group Health.

These two studies of United Healthcare, using different study designs, both reported savings on specialist costs in the UHC gatekeeper plan compared to other plans. However, incentives for gatekeepers alone may not be sufficient to contain overall plan costs. In a subsequent case study of United Healthcare, we examined charges for all enrollees from numerous employers in three states. Although United Healthcare may have contained specialist charges, from 1979 to 1982 specialist charges as a percentage of total charges rose, whereas the percentage of charges by gatekeepers fell.[13]

After operating for nine years in Washington, California, and Utah, United Healthcare was discontinued by SAFECO because of large financial losses, due to management problems and inadequate incentives for specialists and consumers to control costs.[14] The evidence presented in this study suggests that the gatekeeper concept actually may have reduced UHC losses.

Hillman describes the conflict of interest that may exist between financial incentives and providing high quality care.[2] Whenever utilization decreases, especially for referral care, the question of underutilization of services arises. There is no evidence from our studies or those of others which shows plans with gatekeepers providing higher or lower quality care than other plans. Our study was too small to observe any differences in the overall quality of care.

The results of this study are particularly relevant to independent practice associations with incentives similar to United Healthcare. (Since our study population represents employees and their families, these results may not apply to the poor or to retirees.) As the number of gatekeeper plans in the United States continues to grow, the findings of this study have important implications for the design of those plans. First, the United Healthcare incentives for channeling enrollees to primary care physicians increased enrollees' initial use of PCPs and decreased the use of specialists. Second, relatively small financial incentives were sufficient to decrease primary care physician referrals to specialists. It should be emphasized that these incentives applied only to a small proportion of a primarily fee-for-service based practice. Third, the finding of no difference between plans in specialists' use of procedures suggests that in the future plans should incorporate incentives for cost control by specialists.

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Increased Suppressor T Cells in Probable Transmitters of Human Immunodeficiency Virus Infection

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Abstract: To evaluate behavioral and immunologic factors related to transmission of human immunodeficiency virus (HIV) by homosexual intercourse, we studied a population of 329 homosexual/bisexual men (155 partner-pairs) seen in a community health center and medical outpatient clinic. Logistic regression analysis showed that behavioral risk factors for infection in the 130 HIV-infected men included: receptive anal intercourse (OR 4.6, 95% CI-1.8, 12.1); receptive fisting (OR 2.5, CI-1.1, 7.0); nitrite use (OR 2.3, CI-1.2, 4.6); history of gonorrhea or syphilis (OR 2.3, CI-1.4, 3.9); and history of sexual contact with men from areas with many AIDS cases (OR 1.9, CI-1.0, 3.5). Comparing seven men who were probable transmitters of HIV and 11 men who had not transmitted HIV to their uninfected partners despite unprotected insertive anal intercourse, we found no differences in HIV isolation from peripheral blood mononuclear cells, circulating HIV antigen detection, or presence of neutralizing antibody to HIV. Helper T-cell numbers were not significantly different between the two groups, but transmitters had more suppressor T-cells than did nontransmitters. (*Am J Public Health* 1989; 79:1638–1642.)

Introduction

The major behavioral risk factors in the transmission of human immunodeficiency virus (HIV) in homosexual/bisexual men include receptive anogenital sexual contact and a large number of sex partners.[1-3] However, some men with many receptive anal sexual contacts remain uninfected while others with infrequent contact become infected.[4,5] Thus, while particular sexual behaviors may increase the likelihood of HIV transmission, other factors presumably contribute to the efficiency of sexual transmission of the virus.

Transmission of HIV requires that the infected partner transmit the virus and that the exposed partner be susceptible to the infection. Studies of susceptibility to HIV infection of exposed persons have failed to reveal either immunologic or genetic susceptibility factors.[6,7] However, several studies have shown a correlation between worsening clinical status or declining numbers of helper T cells in the HIV-infected partner and increased transmission of HIV.[8-12]

In this study, we evaluated the role of specific sex practices in HIV transmission among homosexual men. We also examined the relation of immunologic laboratory parameters to transmission by comparing seropositive men who appeared to have infected their partners with seropositive men who did not infect their partners despite unprotected insertive anal intercourse.

Methods

Selection of Study Participants

Men were eligible for enrollment in our study if they were homosexual or bisexual, were 18 years of age or older, and resided within the Boston, Massachusetts Standard Metropolitan Statistical Area. Men were seen at the Fenway Community Health Center (FCHC) and/or the New England Deaconess Hospital. AIDS was diagnosed according to the Centers for Disease Control (CDC) classification system (Group IV subgroups C-1 and D-2); AIDS-related complex (ARC) was defined as CDC classification group III (persistent generalized lymphadenopathy), IV-A (fever, weight loss or diarrhea), or C-2.[13] Asymptomatic men were those without AIDS or ARC. Sexual contact was defined as orogenital, anogenital, or oroanal contact involving an exchange of body fluids and/or penetration.

We recruited asymptomatic participants from: men who had at least one visit to the FCHC from May 1984 through March 1987; men referred by friends or physicians outside of the FCHC; advertisements soliciting participation in the study placed in two newspapers; and men referred by persons already enrolled in the study. The protocol for this study was approved by the Human Studies Committees of all participating institutions.

Data Collection

At enrollment, all participants were interviewed to obtain data on demographic characteristics, occupational and travel histories, sexual behaviors, and drug and alcohol use. A complete medical history was obtained for the five years before study enrollment, and a standard clinical examination was performed. All participants were counseled regarding high- and low-risk sex practices and ways to reduce their risk for HIV infection.

Laboratory Studies

All participants were evaluated with a complete blood count (CBC) and lymphocyte subset determination, performed using specific monoclonal antibodies (Ortho Diagnostics, Raritan, NJ). Antibodies to HIV were determined by an enzyme immunoassay (EIA) (Abbott Laboratories, Chicago, IL). All serum samples initially positive by EIA and selected other serum samples (as specified later) were examined by Western blot using previously described methods.[14] Culture of HIV was performed in H9 cells, and reverse transcriptase activity was measured by standard techniques.[15] Antigen-capture assay for HIV employed a sandwich EIA using polystyrene beads coated with human anti-HIV (Abbott Laboratories, Chicago, IL).[16] Neutralizing antibody assays were performed in H9 cells with the HTLV-IIIB isolate of HIV using a modification of the method of

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Robert-Guroff.[17] The assay was performed by mixing 100 50 percent tissue culture infectious doses of HIV with 200 ml of heat-inactivated human serum at various dilutions, and incubating at 4°C for 1.5 hours. Virus was then absorbed onto 1×10^6 H9 cells for one hour and incubated at 37°C. Neutralization titer was defined as positive when the assay blocked 90 percent or more of the control virus infection at a 1:20 dilution.

Statistical Analysis

Confidence intervals for infection rates were calculated based on the chi-square function method of Miettinen.[18] Exposure variables were categorized depending on the distribution of the variable without regard to serologic status. Unadjusted prevalence odds (exposure odds) ratios were estimated and exact 95% confidence intervals (CI) calculated by the method of Rothman.[19] Adjusted estimates were maximum likelihood estimates from logistic regression and 95% CI were calculated using the standard errors of the beta estimates.[20] Continuous laboratory variables were compared using the Wilcoxon 2-sample test.[20] Dichotomous laboratory variables were compared by Fisher's exact test.[20]

Results

All interview information and laboratory samples were obtained at the initial enrollment visit. Three hundred twenty-nine men were enrolled in the study: 23 men had AIDS, 31 had ARC, and 275 were asymptomatic. Six participants were linked to three partners, 12 to two partners, 113 to one partner, and 43 to no partner, for a total of 155 partner-pairs.

The mean age of the study participants was 32.8 years; mean level of education 16 years; median income $10,000–$19,999; and mean length of time residing in metropolitan Boston 144 months. Ten percent of the men had been married; 316 (96%) were White, 9 (3%) were Black, and four (1%) were Hispanic.

One hundred thirty (40%) men were HIV-seropositive, while 199 (60%) were seronegative. Of the 155 partner-pairs, 62 (40%) were concordant seronegatives, 35 (23%) were concordant seropositives, and 58 (37%) were discordant (one seropositive and one seronegative partner). Thirteen of 24 partners (54%) of men with AIDS were seropositive (95% CI 34, 74); 12 of 31 partners (39%) of men with ARC were seropositive (CI 22, 57); 10 of 24 partners (42%) of asymptomatic seropositive men were seropositive (CI 23, 61).

Sexual Practices

To determine which potential risk factors were associated with seropositivity, we analyzed all 329 men in the cohort. In bivariate analysis, six variables were associated with antibody to HIV: number of sex partners in the previous five years; receptive anal sexual activity; receptive oral sexual activity; receptive fisting; sexual contact with persons from areas of high HIV prevalence; history of gonorrhea or syphilis; use of nitrites; and use of other illicit drugs (Table 1).

TABLE 1—Distribution (%), Crude and Adjusted Relative Odds for HIV Infection by Behavioral Risk Factors

	HIV-Antibody Status			Adjusted[a]	
History over Past Five Years	Percent Seropositive (N = 130)	Percent Seronegative (N = 199)	Unadjusted OR	OR	95% CI
No. of male partners					
≤30	36.1	60.4	1.00	—	—
>30	63.9	39.6	2.70	1.36	(.58, 3.22)
No. of men with whom the subject had a specific sex practice					
Receptive penile oral					
1–10	24.2	44.8	1.00	—	—
>10	75.8	55.2	2.53	.83	(.56, 1.23)
Receptive penile anal					
0	6.1	18.8	1.00	—	—
1–10	34.1	57.4	1.84	1.50	(.62, 3.67)
>10	59.9	23.8	7.82	4.63	(1.78, 12.06)
Receptive fisting					
never	87.1	96.5	1.00	—	—
ever	12.9	3.5	4.12	2.54	(1.08, 7.03)
Sex with partners from areas of higher HIV prevalence[b]					
never	20.3	44.6	1.00	—	—
ever	79.7	55.5	3.16	1.88	(1.,01, 3.50)
History of gonorrhea or syphilis					
never	42.1	72.3	1.00	—	—
ever	57.9	27.8	3.59	2.30	(1.36, 3.88)
Nitrite use					
never	12.0	34.7	1.00	—	—
ever	88.0	65.3	3.88	2.31	(1.17, 4.60)
Illicit drug use[c]					
never	8.3	18.3	1.00	—	—
ever	91.7	81.7	2.49	1.37	(.59, 3.22)

a) Adjusted using a logistic regression model that included age and the variables listed in this table.
b) New York City, San Francisco, Los Angeles, and/or Miami.
c) Marijuana, cocaine, PCP, LSD, and/or quaaludes.

A logistic regression model was developed to assess potential confounding by any of the variables significantly associated with HIV antibody in the bivariate analysis. Factors that correlated with seropositivity in the multiple regression model were: receptive anal sexual activity; receptive fisting; nitrite use; history of gonorrhea or syphilis; and sexual contact with persons from areas of high HIV prevalence (Table 1).

Characteristics of Transmitters

We examined partner-pairs in which both partners were seropositive to identify possible transmitters. Ten partner-pairs were identified in which one partner had had over 80 percent of his receptive anal contacts within the relationship (in no case did both partners have over 80 percent of their receptive anal contacts with each other). The partner whose contacts were in large part within the relationship was therefore presumed to have been infected by the other (more sexually active) partner, who was designated the likely transmitter. Peripheral blood cells were available from seven of the transmitters; of these, four had a partner who had no receptive anal intercourse with any other partner, while the other three had partners with 4, 13, and 14 episodes outside the relationship.

Partner-pairs in which one partner was seropositive and one was seronegative were examined to identify men who had exposed their partner to HIV without transmitting infection. All seronegative men were confirmed to be uninfected by negative Western blot and HIV antigen assay. Since we could not be certain that the seropositive partner was positive more than a short time before entry into the study, we excluded partner-pairs who had not had at least six episodes in which the seronegative partner had receptive anal intercourse with an insertive seropositive parther during the six-months before entry into the study. The HIV seropositive partners in these partner-pairs were designated the likely nontransmitters. Sixteen nontransmitters were identified. Peripheral blood cells were available from 11 of these men.

One of the transmitters had AIDS and three had ARC; no nontransmitters had AIDS but three had ARC. Transmitters had longer relationships with their partner (mean ± SD: 58 ± 24 months) than did nontransmitters (36 ± 36 months), but transmitters had fewer total episodes of insertive anal intercourse with their partner (151 ± 185 episodes) than nontransmitters (192 ± 192 episodes). Neither difference was statistically significant. No differences were noted between transmitters and nontransmitters in history of syphilis, gonorrhea, hepatitis B, herpes simplex, herpes zoster, or cytomegalovirus infection, or infectious mononucleosis. Furthermore, transmitters and nontransmitters had similar histories of fever, skin rash, and enlarged lymph nodes.

Transmitters did not differ from nontransmitters with respect to circulating HIV antigen, presence of neutralizing antibody to HIV, or ability to recover virus from peripheral blood mononuclear cells (Table 2). Phenotyping of peripheral blood mononuclear cells of transmitters revealed a median of 1150 total T cells, 430 helper T cells, 960 suppressor T cells, and a median helper/suppressor ratio of 0.40; nontransmitters had a median of 1170 total T cells, 390 helper T cells, 650 suppressor T cells, and a median helper/suppressor ratio of 0.80 (Figure 1). While total and helper T-cell numbers were similar in both groups, transmitters had higher suppressor T-cell numbers than did nontransmitters (Wilcoxon, p = < .02). Helper/suppressor ratios were lower in transmitters than in nontransmitters (Wilcoxon, p = .12). Five of the six transmitters with elevated suppressor cells were re-evaluated after 12 months; only one had persistently elevated suppressor cells.

TABLE 2—Distribution (%) of Laboratory Parameters in Transmitters and Nontransmitters

Parameter	Transmitters (%)	Nontransmitters (%)
HIV culture	1/5 (20)	3/9 (33)
Circulating HIV antigen	1/7 (20)	4/13 (31)
Neutralizing antibody to HIV	3/3 (100)	3/7 (43)

Discussion

Several studies have shown that the major risk factors for acquisition of infection among homosexual men are exposure to a large number of sex partners and exposure to infected persons by receptive anal intercourse.[1-3] Demographic and behavioral analysis of our cohort affirms that it is similar to previously reported cohorts. We assessed risk by estimating exposure odds ratios for HIV infection. The interpretation of such ratios can be difficult when both the temporal sequence and the duration of exposure are unknown. In some cases, the relevant period of risk may have been overestimated, resulting in exposure odds ratios that underestimate the true values.

The risk of acquiring HIV infection was positively correlated with increased number of sex partners with whom the person was the receptive anal sex partner. In our analysis, both fisting and use of nitrites were also associated with seropositivity after controlling for confounding. Fisting might facilitate HIV transmission by causing rectal trauma, thereby allowing virus to bypass mucosal defenses. The biologic mechanism for the increased risk in nitrite users is unclear; the role of nitrites as potential effect modifiers needs to be assessed further.

While these practices are associated with an increased risk of becoming infected with HIV, they do not assure it. Many persons appear to be sexually exposed to HIV on multiple occasions, yet do not become infected.[4,5] Possible explanations for this phenomenon are that some persons are more efficient transmitters than others; that some persons, when exposed, are more susceptible than others; or both.

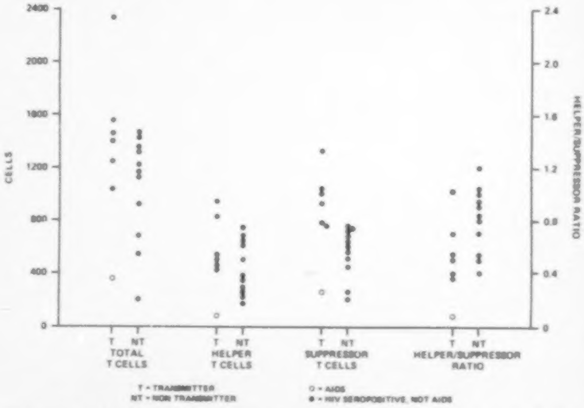

FIGURE 1—T-cell Phenotypic Markers in HIV Transmitters and Nontransmitters; Transmitters had Higher Numbers of Suppressor T-cells

Since previous studies have failed to identify genetic or immunologic susceptibility factors for HIV infection,[6,7] we focused our investigation on factors relating to transmission.

Partner-pairs in which both partners were seropositive were analyzed to select a group of men who were most likely to have transmitted infection to their partners. Such a determination would ideally require that a seropositive man have only been exposed to a single seropositive partner during the time the virus could be assumed to have been circulating in the community. However, monogamous relationships were rare in our cohort. We therefore computed a scale of relative fidelity for each partner. Seven partner-pairs included a man who had had more than 80 percent of his receptive anal contacts since the end of 1981 with one seropositive person (four had 100 percent of their contacts with the presumed transmitter). Therefore, although the date of HIV transmission could not be definitively established and contact with other seropositive partners could not always be completely excluded, we feel the anal insertive partners in these relationships were likely HIV transmitters.

Because of the retrospective nature of the study, the dates of infection of the possible transmitters and nontransmitters were unknown and the exact duration of exposure could not be established. However, the duration of relationships was not significantly different between transmitters and nontransmitters. Moreover, similar numbers of potentially infective episodes were reported for both transmitters and nontransmitters.

We found increased suppressor cells in HIV transmitters. This suggests a role for host immunity in modulating infectiousness by controlling the amount of virus present in host secretions and thereby regulating the dose of virus to which a sex partner is exposed. We were not able to measure the amount of virus in the semen of our transmitters and nontransmitters. However, no relationship was seen between the ability to culture virus from the blood, the presence of neutralizing antibody, or the presence of p24 antigen and the ability to transmit HIV. This may have been due to the small number of samples available for study; alternatively, the presence of detectable HIV in blood may not be related to the presence of infectious virus in semen.

Increased suppressor T-cell numbers have been reported in primary HIV infection.[21] The transmitters in our study therefore may have recently experienced acute HIV infection. Alternatively, since many infectious processes (notably herpes virus infections) can temporarily increase suppressor cells,[22,23] increased suppressor cells in our transmitters may be a marker of an intercurrent infectious process other than HIV infection. Such a process might increase HIV replication by activating HIV amplification genes[24,25] or by interfering with the as yet unknown mechanisms that permit the host to at least temporarily control HIV replication.

Our results suggest that some men with higher numbers of suppressor cells may be more infectious to their sex partners than other infected men. Francis, *et al*, have also reported increased suppressor cells in mothers who appeared to have transmitted HIV to their neonates.[26] On the other hand, several studies have found that HIV-infected men with low helper T-cell numbers or advanced HIV infection are more likely to transmit HIV to their female partners.[8–12] Our HIV-infected subjects were largely asymptomatic with relatively normal numbers of helper T cells; thus we would not have expected to detect transmission by many men with AIDS or low numbers of T-helper cells. Our observations thus support the hypothesis that HIV infectivity is not limited to the period of immunodeficiency seen with advanced HIV infection.[27]

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1987 National Medical Expenditure Survey Data Available on Magnetic Tape

The 1987 National Medical Expenditure Survey (NMES), a nationwide survey sponsored by the National Center for Health Services Research, is available from the National Technical Information Service (NTIS) on magnetic tape. The Institutional Population Component was designed to yield national estimates of the use of, and expenditures for, care during 1987 by persons who reside in nursing homes or facilities for the mentally retarded.

The files contain data from two Phase I instruments: the Facility Questionnaire and the Baseline Questionnaire.

- The Facility Questionnaire was collected from approximately 1400 participating facilities, and the data records contain bed size, type of ownership, facility certification, services routinely provided, average cost, and other information.
- The Baseline Questionnaire file contains data collected from approximately 6,900 persons who were residing in the sampled facilities, and include sociodemographic information such as age, race, sex, information on residences prior to admission, measures of functional limitations and chronic condition, and the work history of persons in facilities for the mentally retarded.

Documentation is provided with the tape and includes a description of NMES, questionnaires, and technical information on sample design, weights, and variance estimation. The NMES is available in EBCDIC or SAS format, and is contained on 9 track, half-inch tape only. Specify density requirement (1600 or 6250 bpi) when ordering.

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4T

A Survival Analysis of Hospitalization among Patients with Acquired Immunodeficiency Syndrome

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AND JOSEPH V. SELBY, MD, MPH

Abstract: Survival analysis techniques were used in estimating lifetime inpatient utilization among patients diagnosed with acquired immunodeficiency syndrome (AIDS) using data on 863 members of the Kaiser Permanente Medical Care Program in the Northern California Region diagnosed with AIDS between January 1, 1981 and June 30, 1987. Using information on both deceased and living patients, we estimated means of 40.3 lifetime inpatient days and 3.3 hospitalizations among all AIDS patients. Those presenting with Kaposi's sarcoma experienced a mean of 7.6 fewer lifetime inpatient days than those presenting with *Pneumocystis carinii* pneumonia (95% confidence interval = .61, 14.6) and a mean of 11.0 (3.9, 18.6) fewer inpatient days than all other AIDS patients. Older patients had fewer hospital admissions than younger ones. Year of diagnosis does not appear to be related to lifetime utilization, and there is an indication that increased survival has been accompanied by decreased inpatient utilization intensity as measured on a per person-year basis. We recommend the use of survival analysis methods in the study of utilization among groups of patients with incomplete follow-up. (*Am J Public Health* 1989; 79:1643–1647.)

Introduction

The cost of providing health care for patients with acquired immunodeficiency syndrome (AIDS) continues to be of national concern. Recent research on the economic impact of the epidemic has provided different estimates of the mean lifetime cost per AIDS patient, ranging from $23,000 to $168,000.[1–6] Much of the variability can be attributed to differences in estimates of the mean lifetime inpatient days accrued by AIDS patients.[1] The disparity in these estimates of inpatient utilization is due in part to true differences across the populations from which the samples of patients were drawn but also in part to the different sampling and estimation techniques. Some investigators have restricted the sample to patients who have died during a given time period, which generally results in a sample biased toward shorter lived individuals. Using survival analysis techniques, we provide estimates of the distribution of lifetime hospitalizations and inpatient days in a large well-defined sample of both deceased and living patients with AIDS.

Methods

Data Collection

We identified 863 cases of AIDS diagnosed between January 1, 1981 and June 30, 1987 who were members of the Kaiser Permanente Medical Care Program in the Northern California Region (KPNCR), a private nonprofit prepaid group practice program with over two million members receiving care in 11 counties; these include the San Francisco Bay Area, a region which has experienced a relatively high incidence of AIDS. According to the California State Health Department, as of June 30, 1987, 4,251 cases were reported in the geographic area served by the Health Plan; we estimate that over 20 percent of these cases were members of the Kaiser Permanente Medical Care Program.

AIDS cases were identified retrospectively by obtaining reports from infection control nurses, and by examining three computerized datasets:

- the Inpatient Utilization System file which contains a record for each hospitalization at any Kaiser Foundation hospital including up to 13 diagnostic codes (*International Classification of Diseases, 9th Revision, Clinical Modification* (ICD-9)) per admission;
- pathology files containing biopsy reports from five major KPNCR hospital pathology departments; and
- a file maintained by the Health Plan's regional pharmacy administration containing all patients who had received zidovudine (AZT).

Each hospital admission from January 1, 1981 through June 30, 1987 was scanned for ICD-9 codes used specifically for AIDS: 279.10 and 279.19, which were used prior to October 1986, and 042.0, 042.1, 042.2, and 042.9 which were used subsequently. In addition, possible cases were identified by searching the hospitalization file diagnosis codes for AIDS-related conditions included in the Centers for Disease Control (CDC) case definition of AIDS.[7] Possible cases were also identified by searching the pathology files for diagnoses of Kaposi's sarcoma, *Pneumocystis carinii* pneumonia, candidiasis, and certain non-Hodgkin's lymphomas indicative of AIDS. For cases identified through the pathology file only or the AZT file only, the medical records were reviewed to verify the presence of AIDS. The medical records of those on the hospitalization file with a possible AIDS diagnosis were also reviewed. Patients were removed from the sample if the CDC criteria were not met.

Analyses are based on the 863 AIDS cases confirmed by chart review (N = 166), infection control nurse reports (N = 171), hospitalization records with diagnosis codes for AIDS (N = 506), and those on the AZT file with an additional record of an AIDS-related infection or neoplasm on the hospitalization or pathology files (N = 5). Males less than 60 years old with a biopsy report of Kaposi's sarcoma and no other source of confirmation were also included (N = 15).

For each case, date of diagnosis and diseases present on that date were obtained from the source of diagnosis confirmation. For those not confirmed either by chart review or nurse report, the earliest biopsy report or hospitalization record with an AIDS-related ICD-9 code was used. Sex and date of birth were obtained from computerized Health Plan

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membership files. It is unlikely that there are patients in this case file who do not have AIDS, but it is possible that some Health Plan members diagnosed with AIDS are not included. Given the variety of sources used to identify AIDS patients, however, the number of unidentified cases is probably small.

The hospitalization file was used to obtain for each of the 863 cases the number of hospitalizations and total inpatient days from date of diagnosis to either date of death, date of leaving the Health Plan, or June 30, 1987, the end of this study. Date of death was obtained by a linkage to death records using the California Automated Mortality Linkage System.[8] Health Plan membership files were used to identify either date of leaving the plan or continued membership in the plan until death or June 30, 1987.

Data Analysis

While survival analysis techniques are often used to study time to an event such as death or onset of disease, these statistical methods are also appropriate for analyzing other random variables.[9,10] One of the main features of these methods is that the random variable of interest can be censored, which occurs when follow-up is incomplete. Information is provided by a censored observation because the final outcome will be at least as great as the measurement taken when follow-up is halted. In this study, a patient's lifetime number of hospitalizations and inpatient days were censored by leaving the Health Plan (N = 13) or by surviving through June 1987 and remaining in the plan (N = 348). We considered a patient's complete lifetime utilization to be measured only if death in the health plan occurred prior to June 30, 1987. The application of these statistical methods to the analysis of inpatient utilization is analogous to the analysis of survival time; we are treating inpatient utilization just as one would treat survival time or any other random variable subject to censoring.

The distributions of survival time, lifetime hospitalizations, and inpatient days were estimated by using the product-limit (Kaplan-Meier) method[11]; associated 95% confidence intervals were formed using Greenwood's estimation method.[12] This method provided estimates of the mean, median, and other percentiles of the distributions of inpatient utilization and survival time. Because lifetime hospitalizations as opposed to lifetime hospital days takes one of a very limited number of values, the sample median and other percentiles are less informative than the proportion of patients attaining selected values, and we present product-limit estimates of the latter. The log rank test[13] was used to assess differences in inpatient utilization distributions across subgroups defined by age, diagnosis at presentation and diagnosis date.

As an alternative to measuring hospital utilization from diagnosis to death of an AIDS patient, utilization rates on a per person-year basis and average length of stay estimates are also presented. Hospitalization and inpatient day rates are calculated for each time period as the observed number of events among AIDS patients divided by the total follow-up time for the cohort in that period. When study subjects enter and exit at any point during the year, this rate is a more appropriate measure of annual utilization intensity than statistics based on annualized data (e.g., multiplying by 12 the utilization by a patient observed for one month). Average length of stay is calculated as the mean number of inpatient days among all admissions during the given time period.

SAS was used for all data analyses[14]; the LIFETEST procedure was used to obtain product-limit estimates of distribution functions, means and percentiles, and to perform log rank tests.

Results

Patient characteristics are listed in Table 1. The mean age at diagnosis was 39.3 years; 98.7 percent were male. Fifty percent of the cases presented with *P. carinii* pneumonia only, and 16.1 percent presented with Kaposi's sarcoma only. Other opportunistic infections accounted for 17.7 percent of the cases. The sex and diagnosis at distributions are similar to those for cases in the same geographic area reported to the California State Health Department except there appears to be a somewhat smaller proportion of KPNCR patients presenting with Kaposi's sarcoma.[6] This may be an artifact due to our reliance on inpatient discharge data for diagnosis information for the majority of cases. In addition, our cases were slightly older than the AIDS patients reported to the State, whose mean age was 38.1 years.[6]

Among all 863 cases of AIDS, the median survival time was 309 days; the 25th and 75th percentiles were 103 and 552 days, respectively. The estimated probabilities of surviving one and two years (and associated standard errors) were .44 (.020) and .20 (.020). The product-limit estimates of the survival distributions for those with an initial diagnosis of Kaposi's sarcoma only, *P. carinii* only, and all others are presented in Figure 1. Diagnosis at presentation was strongly related to survival time (P = .0001). The median time to death for patients with Kaposi's sarcoma was 531 days as compared to 331 days for those with *P. carinii*. The group of AIDS patients with other diagnoses generally had the shortest survival times with a median of 187 days.

The product-limit estimate of the distributions of lifetime hospitalizations and inpatient days are presented in Figures 2 and 3. Only .5 percent of these patients had no hospitalizations. Although the majority of patients (66 percent) were hospitalized for a total of 45 days or less, there was a subset

TABLE 1—Characteristics of 863 AIDS Patients Identified at KPNCR from January 1, 1981 through June 30, 1987

	N	%
Age at Diagnosis (years)		
0–19	4	.5
20–29	110	12.7
30–39	361	41.8
40–49	262	30.4
50–59	97	11.2
60+	29	3.4
Sex		
Male	852	98.7
Female	11	1.3
Date of Diagnosis		
1/81–12/81	3	.3
1/82–12/82	16	1.9
1/83–12/83	39	4.5
1/84–12/84	119	13.8
1/85–12/85	190	22.0
1/86–12/86	282	32.5
1/87–6/87	214	24.9
Diagnosis at Presentation		
Kaposi's sarcoma (KS)	139	16.1
Pneumocystis carinii pneumonia (PCP)	432	50.1
KS and PCP	36	4.2
Other neoplasms	11	1.3
Other infections	153	17.7
Unspecified	92	10.7
Deaths through June 1987	504	58.4

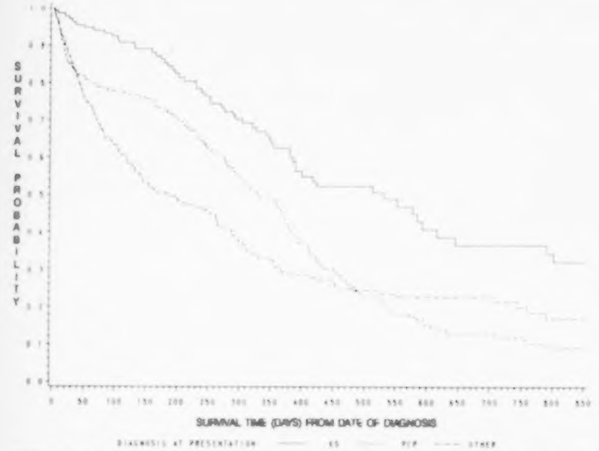

FIGURE 1—Product-Limit Estimate of the Distribution of Survival Time from Date of Diagnosis Among 863 KPNCR AIDS Patients, by Diagnosis (to provide greater detail where most of the sample falls, the distributions are truncated at 850 days).

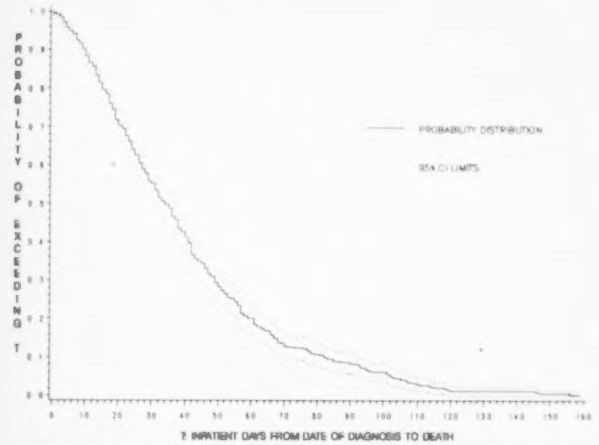

FIGURE 2—Product-Limit Estimate of the Distribution of Lifetime Inpatient Days among 863 KPNCR AIDS Patients Diagnosed between January 1981 and June 1987.

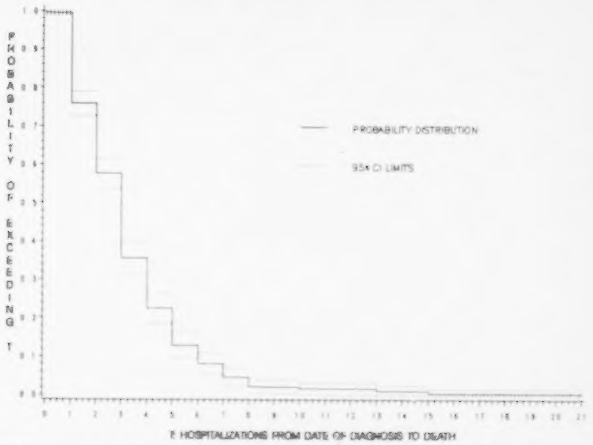

FIGURE 3—Product-Limit Estimate of the Distribution of Lifetime Hospitalizations among 863 KPNCR AIDS Patients Diagnosed between January 1981 and June 1987.

of patients that were relatively high utilizers; an estimated 8 percent were hospitalized more than 90 days. Similarly, 64 percent of the patients required three or fewer hospitalizations, but 8 percent required seven or more.

The distributions of lifetime hospital utilization for all AIDS patients by diagnosis, age, and date of diagnosis are summarized in Tables 2 and 3. The mean number of hospitalizations was 3.3, corresponding to mean and median inpatient days of 40.3 and 34. The distribution of lifetime inpatient days differed among diagnosis groups; those presenting with Kaposi's sarcoma had a lower median number of days compared to those with *P. carinii* and all others. Diagnosis at presentation, however, did not appear to be related to number of hospitalizations. Older AIDS patients experienced fewer hospitalizations but did not differ from younger patients in their inpatient days. Year of diagnosis was not related to either lifetime number of hospitalizations or days hospitalized.

Average length of stay and utilization rates are presented in Table 4. During 1981 through 1983, we observed 105 hospitalizations and 1,767 inpatient days in 32.9 person-years of follow-up on 58 cases, corresponding to utilization rates of 3.2 hospitalizations and 53.7 inpatient days per person-year. Based on 150.1 person-years of follow-up in the first six months of 1987, the hospitalization and inpatient day rates had declined during 1984 through 1986 to 2.6 and 31.1 per person-year, respectively. Similarly, the average length of stay per AIDS-related admission declined from 17.23 days for admissions in 1981, 1982, and 1983 to 10.69 days during the first six months of 1987.

Discussion

Our estimate of mean lifetime inpatient days is substantially lower than the 168 days reported by Hardy, *et al*,[4] and in fact is among the lowest of current estimates as summarized by Bloom and Carliner.[1] Our estimates, however, confirm the finding of other researchers that AIDS patients are hospitalized less often in the San Francisco Bay Area than in other regions of the United States. Scitovsky, *et al*,[3] suggested that strong community-based social support services in San Francisco tend to prevent or shorten hospital stays. They also suggested that in other regions there may be a larger proportion of intravenous drug users, who lack the social support network that other patients with AIDS have developed.

Diagnostic information was obtained from inpatient discharge data for 59 percent of the cases, and thus, there is some misclassification with respect to both date of diagnosis and diseases at presentation. Our estimate of the distribution of survival time from date of diagnosis to death and our utilization analysis by diagnosis subgroup are therefore less reliable than estimates obtained using complete diagnostic information from medical records on all patients. Nevertheless, our survival estimates are in general agreement with other reports from San Francisco.[15] Since it is likely that AIDS patients would generally not be hospitalized without an AIDS-related diagnosis, our primary analyses of inpatient utilization are probably minimally biased by this diagnosis date misclassification.

We examined lifetime hospital utilization among AIDS patients with well-known statistical methods usually used to

TABLE 2—Distribution of Lifetime Inpatient Days among 863 KPNCR AIDS Patients, by Diagnosis at Presentation, Age, and Year of Diagnosis

	N	Median (25th, 75th Percentile)	Mean (S.E.)
Diagnosis			
PCP only	432	35 (17, 56)	39.93 (1.78)
KS only	139	26 (15, 43)	32.33 (3.09)
Other	292	38 (19, 57)	43.38 (2.34)
χ^2 = 8.36, d.f. = 2, P = .015**			
Age at Diagnosis (years)			
0–29	114	38 (16, 68)	47.37 (4.75)
30–39	361	33 (18, 53)	38.83 (1.96)
40–49	262	36 (19, 56)	42.60 (2.59)
50+	126	31 (17, 51)	35.82 (2.54)
χ^2 = 6.18, d.f. = 3, P = .10**			
Data of Diagnosis			
1/81–12/83	58	36 (24, 61)	46.08 (4.76)
1/84–12/84	119	36 (19, 57)	42.71 (2.88)
1/85–12/85	190	36 (19, 53)	39.38 (2.30)
1/86–12/86	282	31 (17, 49)	38.12 (2.48)
1/87–6/87	214	35 (18, 53)	37.73 (4.21)*
χ^2 = 3.85, d.f. = 4, P = .43**			
All cases	863	34 (18, 55)	40.32 (1.32)

*Because the patient with the largest observed number of inpatient days was alive, the product-limit estimate of the mean could not be calculated. The estimate presented is calculated after "uncensoring" this observation and is therefore slightly biased.

**Log rank test for association; d.f. = degrees of freedom.

TABLE 3—Percent Distribution of Lifetime Hospitalizations among 863 KPNCR AIDS Patients by Diagnosis at Presentation, Age, and Year of Diagnosis

		Number of Hospitalizations					
	N	0	1	2	3–4	5+	Mean (S.E.)
Diagnosis							
PCP only	432	0.0	22.8	18.7	32.7	25.8	3.37 (.143)
KS only	139	1.4	21.8	18.8	41.9	16.1	3.11 (.278)
Other	292	0.7	25.2	17.4	36.4	20.3	3.32 (.264)
χ^2 = 1.68, d.f. = 2, P = .43**							
Age at Diagnosis (years)							
0–29	114	0.9	21.6	14.4	35.2	28.0	4.34 (.564)
30–39	361	0.3	22.8	19.2	30.2	27.4	3.26 (.131)
40–49	262	0.0	22.9	18.2	39.9	19.0	3.12 (.152)
50+	126	1.6	28.5	18.6	37.5	13.8	2.93 (.339)
χ^2 = 9.46, d.f. = 3, P = .024**							
Date of Diagnosis							
1/81–12/83	58	1.7	21.8	22.7	37.1	16.6	2.97 (.246)
1/84–12/84	119	0.0	16.5	26.9	34.4	22.2	3.27 (.218)
1/85–12/85	190	1.1	25.9	19.0	32.7	21.3	3.06 (.153)*
1/86–12/86	282	0.4	28.3	11.1	35.3	25.0	3.85 (.455)
1/87–6/87	214	0.0	19.0	17.6	35.2	28.2	3.15 (.221)
χ^2 = 3.40, d.f. = 4, P = .41**							
All Cases	863	0.5	23.6	18.2	35.1	22.6	3.32 (.127)

*Because the patient with the largest observed number of hospitalizations was alive, the product-limit estimate of the mean could not be calculated. The estimate presented is calculated after "uncensoring" this observation and is therefore slightly biased.

**Log rank test for association; d.f. = degrees of freedom.

analyze time to an event. These estimation procedures properly account for the problem of not observing all patients to their deaths. Excluding patients who have not died results in a sample that is biased toward the utilization levels of shorter-lived individuals. Among the patients who died as KPNCR members, the mean number of hospitalizations and inpatient days are 2.6 and 34.6, compared to the product-limit estimates of 3.3 and 40.3; restriction of our sample to the deceased results in a 21 percent reduction in the estimate of mean lifetime number of hospitalizations and a 14 percent reduction in estimated mean inpatient days. Our estimate of the mean lifetime hospital days based on the deceased is extremely close to Scitovsky's[3] estimate of 34.7, which is obtained from a sample of deceased AIDS patients receiving care at San Francisco General Hospital, indicating that the inpatient utilization of the KPNCR patients is similar to that of other AIDS patients in the region.

Our estimates are based on utilization patterns observed only through June 30, 1987. As new therapies are introduced or as patient population characteristics change, our estimates of lifetime utilization may need to be revised. For example, improving survival rates may lead to a change in the inpatient utilization distribution. In this study, there was no evidence of a change in lifetime utilization by year of diagnosis (Tables 2 and 3), although there was a fairly strong downward trend across calendar time in utilization intensity as measured on a person-year basis (Table 4). There was a general increase in survival rates through calendar time. For those with Kaposi's sarcoma, the one year survival rate increased from .60 for those diagnosed in 1981 through 1983 to .65 for those

TABLE 4—Average Length of Stay and Inpatient Utilization Rates among 863 KPNCR AIDS Patients

Admission Date	Average Length of Stay	Observed Person-Years	Number of Hospitalizations* (per person-year)	Inpatient Days** (per person-year)
1/81–12/83	17.23	32.91	3.19	53.69
1/84–12/84	13.78	60.96	3.58	47.68
1/85–12/85	13.56	133.96	3.13	42.67
1/86–12/86	11.72	216.88	2.82	31.43
1/87–6/87	10.69	150.15	2.62	31.10

*Classified according to the date of admission.
**Classified according to the year in which the day occurred.

diagnosed in 1986. For those without Kaposi's sarcoma, the one year survival rate increased from .37 for those diagnosed in 1981 through 1983 to .43 for those diagnosed in 1986. This suggests that the constancy in lifetime utilization while utilization intensity decreased can be explained in part by an increase in survival rates in recent years, i.e., as a group, the same amount of utilization was spread over a longer time period. Also, there appears to be a shift in inpatient utilization from the time near date of diagnosis toward the time of death. Given no change in lifetime utilization, this would result in declining measures of annual utilization intensity during years with less follow-up time close to death than to diagnosis.

Although not presented here, the dependence of utilization on treatment modalities, diagnostic category, and other factors could be examined using methods for regression analysis of survival data.[9,10] In summary, survival analysis methods are appropriate for examining health services utilization in groups of patients with incomplete follow-up. These methods are also useful in estimating the monetary cost of illness with censored data,[6] and an ongoing study of a random sample of KPNCR AIDS patients will examine costs associated with both inpatient and outpatient care.

ACKNOWLEDGMENTS

We gratefully acknowledge the technical assistance of Bill Frank, Merril Jackson, Betty Jue, May Kuwatani, Joanne Owings, Susan Tam, and Joan Thomas.

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Family Access to Health Care: Report Available

Family access to health care is discussed in the report of a conference held in May 1988 at the University of South Carolina and sponsored by the Health Resources and Services Administration.

Single copies of the 136-page report, *Empowering Families for Better Health,* are available from the National Maternal and Child Health Clearinghouse, 38th and R Streets, NW, Washington, DC 20057; tel: 202/625-8410.

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Commentary

A Ten-Year Experience in Control of Poliomyelitis through a Combination of Live and Killed Vaccines in Two Developing Areas

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YITZCHAK SEVER, MD, MICHAEL SCHOENBAUM, BA, AND RACHEL HANDSHER, MSc

Abstract: We describe a successful program of poliomyelitis control using a combination of killed and live polio vaccines over a 10-year period in two developing areas, the West Bank and Gaza, adjacent to a relatively developed country, Israel. During the 1970s, immunization using live trivalent oral polio vaccine (OPV) in these areas covered more than 90 percent of the infant population. Nevertheless, the incidence of paralytic polio continued to be high, with many cases occurring in fully or partially immunized persons. It was thought that this could be due to interference with OPV take by other enteroviruses present in the environment due to poor sanitary conditions in these areas. A new policy combining five doses of OPV with two doses of inactivated polio vaccine (IPV) was adopted and implemented in 1978. In the 10 years since then, immunization coverage of infants increased to an estimated 95 percent and paralytic poliomyelitis has been controlled, despite exposure to wild poliovirus from neighboring countries including an outbreak in Israel in 1988. This experience suggests that wide coverage using the combination of IPV and OPV is an effective vaccination policy that may make eradication of polio possible even in developing areas. (*Am J Public Health* 1989; 79:1648–1652.)

Introduction

The West Bank is located between Israel and Jordan, and Gaza lies between Israel and Egypt (see map). Israel has governed these areas since June 1967. Prior to that, the West Bank was administered by Jordan, and Gaza was under Egyptian rule. In 1988, the Palestinian population of the West Bank was estimated at 887,000, and that of Gaza was 600,000.[1–4] Population density in the West Bank is about 155 residents per square kilometer, and in Gaza about 1,570.[1–4] Approximately 12 percent of the West Bank population, and 53 percent of the Gaza population, live in refugee camps.[2]

Both the West Bank and Gaza are developing areas. In 1987, per capita gross national product was $2,090 in the West Bank and $1,486 in Gaza (up from $836 and $605, respectively, in 1975). In 1987, the crude birth rate in the West Bank was 41 per 1,000 population, while the Gaza birth rate was 48/1,000. In both areas, about half the population is under age 16.[4] Sanitary conditions have improved steadily over the years but still vary widely, with open sewage drainage remaining a problem in Gaza refugee camps.[4,5]

Persistence of Poliomyelitis

In the early 1970s, the Government Health Services in both areas established an expanded childhood immunization program as a major priority. This included four feedings of oral polio vaccine (OPV) in the first year of life.[4–9] By 1975, 85 percent of infants were receiving polio vaccinations, and by 1977 coverage had exceeded 90 percent.[5,7,8] Despite this, however, there continued to be many cases of paralytic polio throughout the 1970s, with many of these cases occurring in children who had received at least one, and up to four, doses of OPV. In Gaza in 1976, half of the 77 polio cases had been fully immunized with OPV, as had eight of 13 cases in 1977.[5]

In response to this problem, a review of polio policy was undertaken in 1978 by the Government Health Services. Vaccine failure was first thought to be due to a deficient cold chain, but testing indicated that the vaccine used in the field was adequately potent.[5,8–10] Diarrheal diseases were widespread, and it was then felt that these might be interfering with OPV take.[5,7,8] It was decided to modify the immunization program to include a combination of inactivated polio vaccine (IPV) and OPV for both the West bank and Gaza, in addition to emphasizing increased immunization coverage and cold chain integrity.[4–6]

Combined OPV/IPV Policy

Since 1978, IPV has been given in the West Bank in combination with diphtheria-pertussis-tetanus vaccine (DPT) in two doses, at 3-1/2 months and 5 months, in addition to five feedings of OPV, given at 2, 3-1/2, 5, 6-1/2, and 12 months. In Gaza, OPV (type I only) is given at age 1 month, followed by IPV and OPV given together at 2-1/2 and 4 months, and OPV alone at 5-1/2 and 12 months. During 1987–88, adequate supplies of IPV were unavailable, and OPV only was used.

Immunization has reached 95 percent of the infant population in both areas, as indicated by immunization data and by serosurveys.[4,8–11] Vaccinations are given through clinics, village health rooms, and mobile services by the Government Health Services, and through United Nations Relief and Works Agency (UNRWA) clinics, which provide primary health care to residents of refugee camps in both areas. UNRWA has followed the combined immunization policy in Gaza; in the West Bank, where UNRWA provides care for only 12 percent of the total population, UNRWA gives OPV alone.

Cases and suspected cases of paralytic polio are hospitalized and reported to public health officials. There is a high

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level of awareness of polio in the medical and general community, and cases are likely to be reported.[4–6] All suspected cases are investigated clinically and with serial laboratory tests carried out at the Central Virus Laboratory of the Israel Ministry of Health. This includes serum antibody levels and stool cultures. Confirmation of cases is based on clinical data, stool cultures, and demonstration of antibody levels increasing over time.[5]

Polio Control

Figure 1 shows annual incidence rates over the 1968–88 period. Both the West Bank and Gaza experienced a decline in polio cases and incidence rates since 1980. During the period 1968–77, Gaza averaged 41.5 cases per year (incidence rate of 10.6 cases/100,000 population). A dramatic decline in polio incidence began in 1978: average cases per year dropped to 7.4 and 0.6 in the subsequent five-year periods (incidence rates of 1.6 and 0.1/100,000, respectively), with one case in 1988. In the West Bank, polio cases averaged some 20 cases per year (3.2/100,000) in the period 1968–77, and began to decline in 1980. There were an average of eight cases per year between 1978–82 (1.1/100,000) and then two cases per year in the 1978–82 period (0.3/100,000), with two cases in 1988.

During 1968–82, Israel averaged 13 cases of polio per year (average annual incidence rate of 0.4 cases/100,000 population). This fell to an average of 1.6 cases per year between 1983–87 (0.04/100,000), but there were 15 cases in 1988 (Table 1).

Table 2 shows a summary of various serosurveys for polio antibody levels carried out in the West Bank and Gaza.[6,8–9,12–14] While the serosurveys did not group children by vaccination status, high levels of immunity in the latter years are due primarily to increased vaccination coverage and improved seroconversion rather than to circulation of wild polio virus, since the incidence of polio decreased during this time.

The 1988 Israel Polio Outbreak

Since 1963, Israel's national routine immunization policy for polio has been to vaccinate infants with four feedings of OPV only. The continued occurrence of endemic polio during the late 1970s, mainly concentrated in five of 15 subdistricts, led to the institution since 1978 of annual spring booster campaigns using type I OPV among all children ages three years and younger in selected areas of these subdistricts.[12,13] Total reported coverage of polio immunization in Israel is over 90 percent.[12–14]

In two other subdistricts in Israel (with 6.7 percent of the total population in 1986), a trial program of three doses of IPV only was carried out between 1982–88.[6,12–14] In 1988, an outbreak of 15 cases of type I poliovirus occurred in Israel; 12 of these cases were from one of the subdistricts (Hadera) using IPV alone. These were largely concentrated in areas of this subdistrict that had serious sanitation problems. The other three cases were temporally and virologically associated with the outbreak, but there was no confirmed physical contact with Hadera. Nine of the 15 cases were 15 years of age or older; nine had received at least three doses of OPV in the past. This outbreak has been described in detail.*[15–17]

The West Bank and Gaza are not immediately adjacent to each other. However, the distances between the most populated areas of the West Bank, Gaza, and Israel average under 50 kilometers. There is very extensive contact between the populations of all three areas (Figure 2).

Despite this, the population of the West Bank and Gaza was unaffected in the 1988 Israel outbreak, although one of the two polio cases in the West Bank may have been epidemiologically connected to the Israeli outbreak. The other case was vaccine associated.** The one case in Gaza in 1988 was epidemiologically unconnected. However, in view of the risk of spread of the disease from Israel to the West Bank and Gaza, a program of administering a booster dose of trivalent OPV to everyone aged 0–40 was carried out, in conjunction with a similar campaign in Israel. This program reached approximately 90 percent of the target population in the West Bank and Gaza.***

The OPV/IPV Debate

The debate on polio policy has been primarily between the Sabin[18] and Salk[19,20] schools who advocate, respectively, OPV-only and IPV-only vaccination policies. The 10-year experience of the West Bank and Gaza reported here suggests that a combined IPV/OPV program can result in polio control even in developing areas with significant sanitation problems and with high potential for the import of wild virus from neighboring areas where wild polio virus is still prevalent.[21]

Oral polio vaccine (OPV) is relatively inexpensive and easy to administer. It induces both humoral and intestinal immunity (secretory antibody response). Due to its oral administration and consequent presence in the intestinal system, it diffuses immunity directly through contact with the vaccinee, and indirectly through the environment, especially in areas with relatively unsanitary conditions.[18,22–29] OPV-induced immunity among contacts of the vaccinee and through the environment is a major advantage over using IPV alone.[25,27]

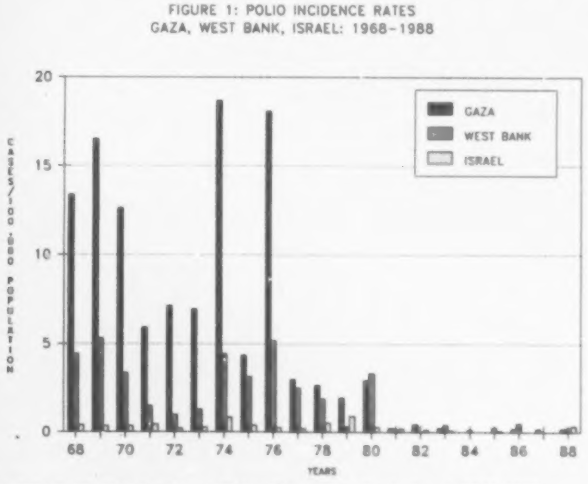

FIGURE 1—Polio Incidence Rates, Gaza, West Bank, Israel: 1968–1988

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*Melnick JL, Orenstein WA, Rey M: Report of a consultation on poliomyelitis control in Israel, October 5–6, 1988. Jerusalem: Ministry of Health, 1988 (unpublished); and Slater PE, personal communication, 1989.

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TABLE 1—Confirmed Poliomyelitis by Five-Year Periods: Cases, Average Annual Cases, and Annual Average Incidence Rates, West Bank, Gaza, and Israel, 1968–88

	1968–72	1973–77	1978–82	1983–87	1988
Gaza					
Cases	205	210	37	3	1
Average annual cases	41.0	42.0	7.4	0.6	—
Average annual incidence rate	11.1	10.2	1.6	0.1	0.2
West Bank					
Cases	92	110	40	10	2
Average annual cases	18.4	22.0	8.0	2.0	—
Average annual incidence rate	3.1	3.3	1.1	0.3	0.2
Israel					
Cases	53	67	77	8	15
Average annual cases	10.6	13.4	15.4	1.6	—
Average annual incidence rate	0.3	0.4	0.4	0.04	0.34

NOTE: Incidence rates are cases per 100,000 population.

TABLE 2—Serosurveys for Percent with Protective Antibody Levels by Polio Type, West Bank and Gaza

Area/Age Group	Year	No.	Type I	Type II	Type III
Gaza					
6–36 months	1974	50	56%	74%	66%
	1978	50	60	82	70
	1980	117	100	98	100
	1983	209	89	83	82
West Bank					
6–8 years	1983	240	92	96	97
7 years	1986	266	97	99	96

NOTES: Serosurveys were performed with 1:4 dilutions (1974 in Gaza used 1:10). Survey samples were random (1980 in Gaza was follow-up survey of vaccinees only).

However, successful immunization with OPV requires four doses at appropriate intervals. In addition, because OPV contains an attenuated live virus, there are rare cases (1 per 2.6 million doses in the United States for 1974–84) of vaccine associated disease, in which the vaccinee or a person in contact with the vaccinee incurs paralytic polio from mutated vaccine virus.[22,25,27,30–31] Additionally, there is strong evidence that OPV take is inhibited if the vaccinee is infected with other enteroviruses, resulting in failure of the vaccine to produce satisfactory levels of protective antibodies.[23,24,26,31]

Inactivated polio vaccine (IPV) rapidly produces a high level of protective antibodies. It poses no risk of vaccine-associated disease, and is not susceptible to interference from other enteroviruses.[26] IPV also protects vaccinees (although not necessarily their contacts) from vaccine-associated disease contracted from OPV doses administered subsequently.[25–27] Newly developed enhanced-potency IPV vaccines are even more effective in terms of speed and level of immunogenicity than previous inactivated vaccines.[22,24]

However, IPV produces only limited intestinal immunity. This permits vaccinees who receive only IPV to be acute carriers of polio. The wild polio virus can enter and infect the intestinal tract of vaccinees, who will not themselves contract paralytic polio, but may excrete the virus, spreading it through the environment or directly to other people.*[25–27]

The Combined OPV/IPV Approach

Both vaccine types have proven effective in controlling

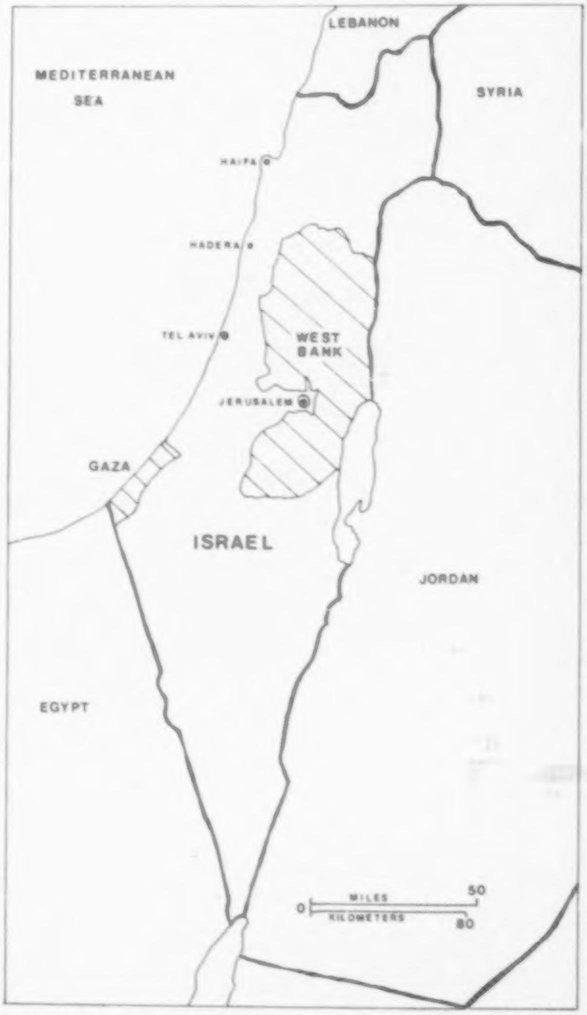

FIGURE 2—Map of Israel Showing West Bank, Gaza and Hadera (site of 1988 outbreak)

polio in developed and developing countries. However, particularly in less than ideal sanitary conditions, neither type used alone seems capable of eradicating the disease. Recently, the combination policy of IPV and OPV administered sequentially has attracted interest internationally.[22,26–27,32–33] Denmark has been using such a policy successfully since 1968.[27,32]

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*Melnick JL, Orenstein WA, Rey M: Op Cit.

Although poliomyelitis is an acute viral disease affecting humans only, with no animal host and no chronic carrier status, it can be imported by infected persons from outside a given geographical area. Polio virus can then be transmitted by contact or through sewage and waste processing systems.

Developing countries often face great difficulties in vaccinating the numbers of people necessary for adequate herd immunity. In addition, they face problems of interference with vaccine take from other enteroviruses. The policy of combining OPV and IPV is well suited for minimizing the effects of these problems.

As the 1988 polio outbreak in Israel illustrates, an IPV-only vaccination policy for areas with sanitary problems, and where wild polio virus is present, has potential for failure. The 1985 polio outbreak in Finland, even though associated with an inadequately immunogenic IPV (for type 3), further suggests that an IPV-only policy may not produce adequate levels of herd immunity even in areas without serious sanitary problems and where wild polio virus introduction may be unlikely.[21,28] In Holland, where IPV has been used alone, repeated introductions of wild polio virus were associated with relatively large numbers of cases (114 over the period 1974–84), mostly among non-immunized members of religious communities. Introduction of wild polio virus to an area by travelers to and from developing areas is a risk factor that must be considered when determining polio vaccine policy.[25]

The spread of immunity through personal and environmental contact from OPV use is a major advantage; the reduction or elimination of vaccine-associated polio by use of IPV is also a major advantage. The West Bank and Gaza experience show the effectiveness of combining IPV with OPV in areas where OPV alone has a high failure rate.

The combined policy has the further advantage of at least partially mitigating the problem of incomplete vaccination schedules. Giving OPV and IPV at the same time makes it likely that infants, particularly in populations where return visits for further vaccinations cannot be guaranteed, will be protected even after only three visits although a complete schedule is more desirable. The new, high potency IPV's ability to rapidly create high levels of protective antibodies makes seroconversion more likely, even with incomplete vaccination schedules. In the applied field setting, the combined policy can lend itself both to routine campaigns and to the blitz campaign approach used in some developing countries.[24,31,34]

In addition, the combined IPV/OPV policy can reduce or eliminate the risk of vaccine-associated polio, which remains an important factor in the epidemiology of the disease. As an example, four of the eight Israeli polio cases occurring between 1982 and 1987 were vaccine-associated, even though the rate of vaccine-associated polio reported in the United States is considerably lower. For this reason, IPV should be given before or with the first OPV dose, since the greatest danger of vaccine-associated disease comes from the first dose of OPV in individuals previously unvaccinated against polio.[29]

International Implications

The global eradication of smallpox was an outstanding achievement, with enormous benefit to the world. It was achieved by a combination of a safe, effective vaccine, new technology in vaccination, a world consensus on the need to act, and excellent strategies in application of the technologies available.[23] As the World Health Organization (WHO) now has targeted polio for eradication, it is vital to develop the most appropriate strategies in the use of existing and new technologies. Polio eradication will be more complex than was smallpox eradication because of differences in transmission, clinical manifestation, interference, and duration of immunity. Perhaps most importantly, immunity to polio requires multiple doses and boosters, whereas a single dose of smallpox vaccine confers lifelong immunity.[23]

As a result of the importance and complexity of the issue, documented field trials of various approaches to polio control are needed. The West Bank and Gaza experience in use of the combined OPV-IPV approach should be considered for replication in wider field trials in the search for appropriate strategies in polio control/eradication, particularly for developing countries. The Israel Ministry of Health has recently decided to adopt the combined OPV/IPV approach for its future immunization policy. Although the combined approach is more expensive than OPV alone, its advantages make the combined approach an essential tool in the world battle to eradicate polio.

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Business and Nursing Schools Join Ranks to Develop MBA-MSN Degree Programs

Ten of the nation's top-rated graduate schools of business are joining ranks with their graduate schools of nursing to provide future nursing leaders with the same caliber of management, financial, and analytic training usually reserved for the business world. Superior management skills—combined with their clinical skills—are essential to tomorrow's nurse leaders. This unprecedented alliance between the schools of business and nursing "to develop joint MBA-MSN degree programs comes in recognition of the enormous resource allocation, quality control, and employment practice decisions now made by nursing leaders," said Dean Meyer Feldbert of Columbia University Graduate School of Business.

Nursing comprises the largest workforce in the health care industry, and health care is the nation's third largest industry, with expenses approaching $2 billion per day.

Almost 300 nurses with MBA degrees were recently surveyed for The Commonwealth Fund by Louis Harris and Associates, which found that most nurse-MBAs get far more managerial responsibilities as a result of their MBA degree. The most valued skills, according to the nurse-MBAs surveyed, were in finance and marketing. Eighty-eight percent of respondents advised other nurses like themselves to get an MBA degree.

Each of the universities developing the new joint degree programs will receive $100,000 in start-up support from The Commonwealth Fund. The Fund will also provide fellowships of $15,000 each to 25 nurses per year, as further encouragement of this type of education.

The 10 universities developing MBA-MSN graduate degree programs are: University of California (Los Angeles), Case Western Reserve University (Cleveland), Columbia University (New York City), Iowa University (Iowa City), University of Michigan (Ann Arbor), University of Rochester (New York State), Rush University/Northwestern University (Chicago), University of Texas (Austin), Vanderbilt University (Nashville), and University of Virginia (Charlottesville). Courses are due to start in Fall 1990.

The Commonwealth Fund is located at Harkness House, One East 75th Street, New York, NY 10021. Tel: 212/535-0400.

A Comparison of Perceived and Objective CVD Risk in a General Population

Minoo Niknian, PhD, Sonja M. McKinlay, PhD, William Rakowski, PhD, and Richard A. Carleton, MD

Abstract: This report examines correspondence between perceived and objective (American Heart Association's RISKO) risk of heart attack and stroke in a randomly selected sample from two surveys (n = 4,171) conducted in each of two New England cities, in 1981–82 and 1983–84, respectively.

Results confirmed prior reports that people tend to underestimate their CVD (cardiovascular disease) risk and showed that estimates of those at lowest risk were most accurate. (*Am J Public Health* 1989; 79:1653–1654.)

Introduction

Self-assessments of health status have been an integral part of behavioral science research and theory. One question of continuing interest has been the correspondence between "perceived' and "objective" indices of current health. As a general rule, self-assessed health appears to have a basis in actual physical status, using professional or laboratory indices of illness severity as a standard for comparison.[1-3] However, it also seems well-accepted that when discrepancies between personal and objective data occur, individuals are more likely to overestimate their health status.[4-6] This report examines the correspondence between perceived and "objective" risk of heart attack or stroke in a representative community sample.

Methods

Data were collected in two community health surveys, conducted in each of two New England cities in 1981–82 and 1983–84, respectively. Each spanned an entire year to control for seasonal differences. Approximately 5,244 interviews with adults 18–65 years were completed and two blood pressure readings, height, weight, total cholesterol and HDL were measured. Response rate averaged 70 percent.[7]

Two variables were used to construct the dependent variable for this analysis: *Perceived Cardiovascular (CVD) risk* and total *RISKO score* (from the American Heart Association's CVD risk assessment instrument).[8] RISKO was calculated with height, weight, blood pressure, cholesterol and estrogen (for women) as obtained directly from the surveys (n = 4,171). A second, estimated score, using the RISKO substitution rule was analyzed, produced equivalent results and is not reported here.*

RISKO ranges from 0 (lowest risk) to 24 (highest risk) for each of four age × gender groups (<55, ≥55 years; male and female). Three risk categories were constructed from this score: "low risk" was defined as having all except one risk factor at the lowest risk level for those aged <55, producing a maximum score of 5. Similarly, the "high risk" category was defined as having all factors at the third risk level or higher in the younger age category, producing a minimum score of 12. The three objective risk categories, defined on RISKO scores were therefore (0–5, 6–11, 12–24). Because RISKO is conditional on age × sex category, a score of 9 has the same relative meaning of average risk, in each age × sex group.

Perceived risk was assessed as: "Compared with persons of your own age and sex, how would you rate your risk of having a heart attack or stroke within the next five years?" The pre-coded responses were: high (3), average (2), low (1), and don't know. Persons responding to this last category (7 percent) were excluded from subsequent analyses.

Independent variables considered in this analysis were: age (in years), gender, years of schooling, ethnicity (Portuguese, French Canadian, and other), body mass index (Wgt [Kg]/hgt^2 [m^2]); smoking status (yes/no); taking any CVD-related medication (yes/no); and knowledge (whether or not at least one CVD risk factor is mentioned as modifiable).

The discrepant categories between RISKO and perceived risk have possible values (−2, −1, 0, 1, 2), where 0 represents agreement. Using this score as the dependent variable, a stepwise regression[9] was performed on all the independent variables, entering sociodemographic characteristics first. Variables were retained in the model if their individual F-values were at approximately the .01 level and/or induced a relative increase in R^2 of greater than 5 percent. An equivalent discriminant analysis, as a validity check of this approach, produced the same models.* The second stage involved calculation of weighted and unweighted Kappa statistics,[10] which measures agreement between two scoring systems, independently applied to the same subjects.

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*Data available on request to author.

TABLE 1—Unadjusted Kappa and other Agreement Statistics for Comparisons of Objective-Perceived Risk*

Comparison	n	$\hat{K}$	S.e.($\hat{K}$)	% Agreed	% Underestimating Actual Risk	% Overestimating Actual Risk
RISKO	4171	.076	.010	42.8	39.7	17.5

*Note: Reference from all analyses for RISKO using the substitution rule are available upon request.

Results

As expected, educational level had a significant impact on congruence, as did age and gender. Interestingly, knowledge of CVD risk had no association with discrepant scores. Only the two well-known CVD risk factors (smoking and BMI) contributed to explaining discrepant scores, as did use of hypertension medication (diuretic or beta-blocker).

To further investigate the concordance between perceived and actual risk, Kappa statistics were calculated[10] for each level of each of the significant covariates mentioned above as well as the percent of responses in agreement. Females (44.4%) were more likely than males (40.5%), while younger people (44.3%) were more likely than those over age 44 (40.3%) to perceive their risk accurately. Education, dichotomized at 16 years, was positively related to higher agreement (47.1% compared to 42.5%). This is consistent with the regression results using the continuous measure of education. Those at lowest CVD risk [non-smokers (44.6% versus 40.0%) and those with BMI ≤22.5 (49.8% versus 40.7%)] perceived their risk more accurately, as would be anticipated. Those on hypertensive medication also perceived their risk more accurately (45.5%) than those not on medication (42.4%).

Apart from these associations, the overall Kappa was .078 which is low and never exceeds 0.12 indicating that agreement was generally poor. The direction of disagreement was further investigated (Table 1) by considering the percentage under- and overestimating their risk. From these statistics it is evident that twice as many people underestimated their risk, compared to those who overestimated their risk.

Discussion

The analyses presented here lead to two major conclusions. First, there was poor agreement between self-perceived and a more objectively assessed cardiovascular risk, with a high prevalence of risk underestimation. Second, higher levels of agreement between perceived and objective risk are associated with being female, being younger, having more education, being a non-smoker or on hypertensive medication, or having a lower than average BMI. As expected, people at low CVD risk are the most likely to correctly define themselves. Hypertensives, as the clear exception, are sensitized to their higher risk. Knowledge of CVD risk was not independently associated with increased accuracy because of its high correlation with educational level.

The results presented here confirm prior reports that people generally underestimate future risk of negative health outcomes. The important additional contributions of this study are:

- Use of a representative, randomly selected community sample of adults; and
- Use of a validated scoring system (RISKO) to assess cardiovascular disease risk within a specific future period (five years).

In this environment of heightened awareness and decreasing mortality rates, the persistence of CVD risk underestimation is of substantial concern. Clearly there is still considerable room for improvement in health education efforts, at least with respect to CVD.

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hourly employees and the start of programs for salaried employees encouraging early retirement and separation from the corporation.[10] The number of inactive employees (retirees and surviving spouses) grew twice as fast as the active workforce. Dependents, who comprise over half of the GM population,[7] accounted for much of the growth. The number of spouses grew for all employee types except active salaried employees, and grew very rapidly, nearly 3.5 percent per year. In contrast, the number of children hardly grew at all.

Discussion

This is the first study to examine demographic change in a population provided health benefits by a major corporation. It has documented rapid growth, faster than in the US, in this large group. If current trends continue, this population will become larger, older, and have a greater percentage of women.

National health care utilization statistics[11-12] suggest that these demographic changes alone increased GM hospital days of care by 10 percent and physician visits by 7 percent from 1983 to 1987. Consequently, it is not surprising that GM's health care costs rose from $2.2 billion in 1983[5] to nearly $3 billion in 1987.[6]

The trends identified in this study may or may not predict future demographic conditions. Future investigations should estimate the impact on growth rates of employment conditions, sociodemographic processes influencing population growth, and administrative changes.

This study has found that despite increased employment, dependents, particularly spouses, were responsible for most of the population growth. This finding demonstrates that much remains to be learned about how individuals, particularly nonworkers, become eligible for employer-sponsored health coverage. Case studies of employers, such as this one, can illuminate findings from general population surveys about the link between employment and health care coverage.[13] In response to potential legislation that increases the cost of health benefits, employers could pay more attention to who is eligible for these benefits in their efforts to provide cost-effective coverage.

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Fatal Unintended Carbon Monoxide Poisoning in West Virginia from Nonvehicular Sources

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Abstract: Based on medical examiner reports and state vital records, 1978–84, nonvehicular carbon monoxide (CO) caused 62 unintended deaths, representing 42 percent of all unintended fatal CO poisonings in West Virginia. Sources were almost always heating or cooking appliances associated with incomplete combustion of fuels (methane, butane, or propane) not commonly recognized for their potential to produce CO. Hazards included failure to provide recommended venting, neglected maintenance, or use in small areas without natural ventilation. (*Am J Public Health* 1989; 79:1656–1658.)

Introduction

Carbon monoxide (CO) is the agent most often responsible for fatal, unintended poisoning in the United States.[1] Motor vehicle exhaust causes the majority of cases, but a substantial number occur in domestic or recreational settings[1] where sources are not adequately identified in national mortality statistics[2] or by death certificate review.[3] While heating appliances are recognized as important sources of CO exposure,[3-9] their overall role and specific hazards have not been carefully examined in a population-based study.

Using data obtained principally from the Office of the Chief Medical Examiner (OCME), we have reviewed all deaths in West Virginia attributed to unintended CO exposure from nonvehicular sources during a seven-year period.

Methods

The West Virginia Medical Examiner has jurisdiction over all homicidal, suicidal, accidental, and unattended

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natural deaths in the state and maintains a centralized case record file. When indicated, analyses for blood alcohol concentration (BAC) and percent carboxyhemoglobin (COHb) are performed by standard methods.[10,11]

We reviewed the files for the period 1978–84 to identify all certified unintentional deaths attributed to CO from nonvehicular sources, excluding uncontrolled fires. We also reviewed vital statistics for in-state deaths coded "accidental poisoning from gases and vapors" during 1978 (ICD-8 codes E870-7) and 1970–84 (ICD-9 codes E867-9) to identify any additional unintended, nonvehicular CO deaths not reported to the OCME. Data recorded included each victim's age, sex, BAC, and percent COHb, along with date and place of death, and the alleged CO source. Hazards identified at the site of exposure were recorded from information either in the OCME record or acquired through special request to local police or fire departments.

Results

From 1978–84, there were 62 unintentional deaths in West Virginia attributed to CO from nonvehicular sources (two were recorded by vital statistics alone). These deaths occurred in 38 domestic episodes (51 deaths) and in seven recreational settings (11 deaths). On the death certificates, 31 (50 percent) fatalities were not specifically attributed to CO and one was mistakenly associated with motor vehicle exhaust (Table 1).

Postmortem COHb was measured in 47 victims and ranged from 50–88 percent. For eight victims who were not tested, death from CO was determined by association with confirmed cases (three deaths), by CO-positive environmental samples (three deaths), or by elevated premortem COHb (two deaths). For seven remaining victims, CO was assumed the cause based on observations at the scene.

Fatal episodes increased sharply in the fall and peaked during the winter (Figure 1). The male to female victim ratio was almost 4:1 (Table 2). Death rates were highest in young adults and the elderly. Blood alcohol was tested in 40 (70 percent) of the 57 victims ≥15 years of age. Nine (22.5 percent) tested positive, with BAC ranging from 2–46 mmol/L; only three (7.5 percent) had levels consistent with legal intoxication (22 mmol/L).

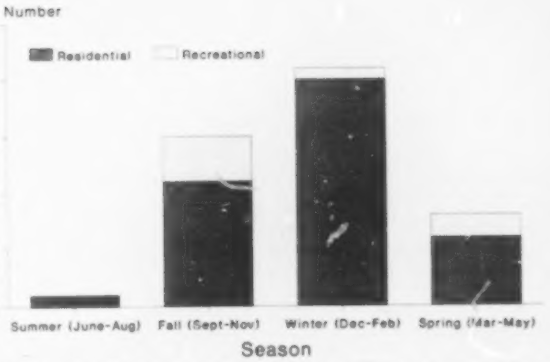

FIGURE 1—Fatal Unintentional Carbon Monoxide Poisonings in Residential and Recreational Settings by Season, West Virginia

TABLE 2—Deaths from Unintentional Nonvehicular CO Exposure by Age and Sex of Victim, West Virginia, 1978–84

	Deaths			
Age (years)	Male	Female	Total	Rate*
0–14	3	2	5	1.6
15–24	13	3	16	6.6
25–44	15	3	18	5.0
45–64	10	3	13	4.7
≥65	8	2	10	6.0
All ages	49	13	62	4.5

*Deaths per 1,000,000 population per year

Residential Settings

Furnaces or space heaters were implicated in 34 (89 percent) of the 38 residential episodes. Cooking ranges were the source of CO in three (8 percent) episodes and a generator was the source in one. Thirty-six units (94 percent) were fueled by natural gas (methane) or butane (only one unit). Table 3 shows that on-site investigations found hazards related to:

- installation of large volume furnaces dependent on direct venting of combustion products to the outside;

TABLE 1—E-Code Classifications on 62 Death Certificates for Unintentional CO Poisoning from Nonvehicular Sources, West Virginia, 1978–84

E-code	Description of Accidental Poisoning	Deaths
ICD-8*		
874	CO from incomplete combustion of domestic fluids	4
875	Other CO	1
ICD-9**		
867	Gas distributed by pipeline	1
868.0	Liquefied petroleum gas in mobile containers	4
868.1	Other and unspecified utility gas	26
868.2	Motor vehicle exhaust gas	1
868.3	CO from incomplete combustion of other domestic fuels	22
868.8	CO from other sources	1
868.9	Unspecified CO	2

*in use for 1978
**in use for 1979–84

TABLE 3—Cause of Hazard and Fuel Responsible for Fatal Unintentional Exposure to CO from Nonvehicular Sources by Setting, West Virginia, 1978–84

Setting and Cause of Hazard	Episodes	Fuel
Residential		
Faulty installation		
Failure to install or connect vent	4	Methane (4)
Vent obstructed during installation	3	Methane (3)
Faulty vent design	3	Methane (3)
Neglected maintenance		
Burner malfunction from accumulated debris or faulty adjustment	5	Methane (4), Butane (1)
Damaged vent*	4	Methane (4)
Use in sealed or confined enclosure	13	Methane (12), Kerosene (1)
Miscellaneous	1	Gasoline (1)
Not identified	5	Methane (5)
Recreational		
Use in confined enclosure	7	Butane (2), Propane (3) Charcoal (1), Wood (1)

*Includes combustion product leakage through loose indoor vent joint (three episodes) or backflow caused by downdraft from corroded exterior vent segment (one episode).

- use of smaller units not requiring direct venting but operated in closed rooms. Windows were sealed with plastic in at least four of these 13 episodes;
- neglected maintenance.

Recreational Settings

Five recreational episodes involved the use of butane or propane-fueled heating stoves in poorly ventilated enclosures (two tents, two recreational vans, and a cellar house); one resulted from a smoldering charcoal grill in a recreational van; and one resulted from a poorly vented wood-burning fire in an enclosed outdoor shelter.

Discussion

Despite the elimination of CO from domestic gas since the 1940s and '50s with conversion to natural gas, and the widespread use of other clean burning fuels, domestic and recreational CO exposure remains an important health hazard in West Virginia, contributing 62 (42 percent) of the 149 unintended CO deaths from 1978 through 1984; 87 deaths from motor vehicle exhaust were described earlier.[12] [illegible] use was not a common predisposing factor in these exposures as it was for exposures to motor vehicle exhaust. For surveillance purposes, coded death certificates alone grossly underestimate the problem in nonvehicular settings.

While each of these episodes reflects some failure to adequately remove CO from the living or recreational environment, the vast majority are associated with fuels not commonly recognized for their potential to produce CO.[13] Unlike cruder fuels (gasoline, kerosene, wood, and coal), methane, butane, and propane normally undergo complete combustion in the presence of sufficient oxygen, producing nontoxic carbon dioxide and water vapor.[14] In most instances, CO production by these fuels had at least one obvious explanation:[14]

- operating small volume, unvented gas-burning heaters in closed spaces without natural ventilation (e.g., open windows) for replacing depleted oxygen;
- inadequate venting (removal) of combustion products, displacing oxygen from the ambient supply and/or preventing adequate air flow to the burners. While this was usually associated with faulty installation, it occasionally reflected long-term use and neglected maintenance;
- neglecting to adjust or clean burners, causing inadequate mixing of oxygen and fuel in the combustion chamber.

Prevention therefore depends largely on assuring proper function and appropriate use of gas-burning applicances and can be achieved by professional installation, annual servicing, and education to avoid burning fuels in confined or tightly sealed enclosures. A coordinated educational and inspection program involving health departments, gas companies, and appliance distributors—conducted in the late summer and early fall—would be desirable. Because there is a tendency to conserve heat by closing or sealing windows, terminating the production of unvented units has also been suggested.[7]

Nonlethal CO exposure is a significant health hazard[4,6,7] that can be overlooked when symptoms of headache, nausea, vomiting, and malaise are attributed to other conditions. Several reports of fatalities among persons who had previously consulted a physician[8,9,13] suggest that health care providers can help to identify homes with a CO hazard by recognizing symptoms of exposure.

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2T

Gastrointestinal Effects of Water Reuse for Public Park Irrigation

ROGER DURAND, PHD, AND GERHARD SCHWEBACH, MS, MPA

Abstract: To investigate the gastrointestinal effects of employing recycled water as an irrigation source for urban public parks, we studied subjects active in parks irrigated with potable water, nonpotable water of wastewater origin, and nonpotable water of runoff origin. Wet grass conditions during activity and elevated densities of common indicator bacteria, but not exposure to nonpotable irrigation water per se, were found associated with an increased rate of gastrointestinal illness. (*Am J Public Health* 1989; 79:1659–1660.)

Introduction

Water reuse or recycling has become a widely practiced solution to the water shortage problems of arid and semi-arid regions of the United States.[1-3] To date, however, there is almost no evidence available on the health effects of using recycled water as an irrigation source for public access parks.[4] This paper reports the results of an investigation of the gastrointestinal consequences of employing recycled water for irrigating urban parks. The investigation itself, conducted in the city of Colorado Springs over the period from 1984 to 1987, was designed to answer three research questions:

- Does the irrigation of public access parks with recycled nonpotable, rather than potable, water result in a greater incidence of reported gasrointestinal illness among park users?
- Do higher degrees of exposure to nonpotable water used in irrigation lead to an increased incidence of reported gastrointestinal illness among park users?
- Can an indicator bacteria and bacterial density that is associated with an increased rate of gastrointestinal illness be identified?

Methods

In order to address the research questions, a prospective cohort research design was utilized. Epidemiologic data were collected by means of telephone surveys over two irrigation seasons from randomly selected subjects active in parks irrigated with potable water, nonpotable water of wastewater origin, and nonpotable water of runoff origin. The nonpotable water of wastewater origin utilized for park irrigation was effluent that had been sand-anthracite filtered and chlorinated up to a level of 4–6 mg/L. The nonpotable irrigation water of runoff origin was untreated city runoff that was collected in a pond.

During an initial interview, conducted within two days of park activity, each subject's presence in a park at the scheduled time was confirmed. Each research subject was then asked a battery of questions concerning his/her park activity and exposure to the park's irrigation water. The same subjects were contacted again 10 to 12 days later to determine whether they had experienced an illness in the intervening time period and, if so, the nature of the symptoms they experienced. Following studies sponsored by the Environmental Protection Agency (EPA),[5] "credible symptoms" of gastrointestinal illness were studied as the principal dependent variable. Such symptoms were defined to include reports by subjects of any one of the following accompanying "stomach discomfort": diarrhea, vomiting, cramps, fever, weight loss, excessive gas, or blood in stools.

A total of 2,642 subjects were surveyed on both the initial and follow-up interviews for a completion rate of 94 percent among those potential subjects who could be contacted. Unsuccessful attempts were made to contact an additional 511 subjects. All subjects were screened to avoid including them in the study more than once.

In addition to the survey evidence, microbiological samples of irrigation water were collected periodically in the parks included in the study that were irrigated with nonpotable water. A few samples were also obtained from parks employing potable irrigation water. Analyses were performed on these samples both for common indicator bacteria and for the biological pathogens *Salmonella* and *Shigella* using standard methods of microbial analysis.[6]

Results

Adjusted for the effect of gender, subjects exposed either to wastewater or to run-off water reported credible symptoms of gastrointestinal illness (6 percent and 5 percent, respectively) about as frequently as subjects exposed to potable irrigation water (7 percent) (Table 1). Unadjusted rates for respondents active in parks irrigated with potable water were higher than among counterparts active in parks irrigated with wastewater and runoff at statistically significant levels, but adjustments for variations in the gender and age composition of the parks clientele eliminated this initial difference. Separate controls for type of park activity (soccer and flag football versus golf, softball, and senior citizen games) and for variations in degree of exposure to irrigation water (wet grass, sprayed by a sprinkler, frequency of prior exposure and length of exposure) did not reveal a greater relative incidence of reported symptoms among subjects exposed either to wastewater or to runoff.

Interestingly, wet grass conditions during activity, regardless of type of park irrigation water, were found associated with greater symptoms reporting (adjusted difference = 3%, 95% Cl = 1, −1), a finding consistent with some previous research.[7]

To eliminate the effect of interpark mobility or "park switching," we examined the 105 individuals from parks irrigated with potable water who had participated in an activity in a park irrigated with nonpotable water prior to the follow-up survey. The reported rate of illness was the same as in the remaining 966 subjects who did not switch.

Subjects active in parks irrigated with runoff water were found to be exposed to higher average (mean) densities of fecal streptococci, fecal coliforms, and total coliforms than were other subjects. The mean fecal streptococcus and the

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TABLE 1—Reported Symptoms by Park Irrigation Type Adjusting for Demographic Characteristics and Exposure

Park Type/ Adjustment for:	Cases	Total Subjects	Adjusted* Reported Cases (%)	Difference in Adjusted Percentage	95% Confidence Interval
Potable Water					
Unadjusted	90	1071	8	—	—
Gender	90	1071	7	—	—
Age	90	1062	8	—	—
Park activity	90	1071	9	—	—
Wet grass	90	1071	9	—	—
Sprayed	90	1067	8	—	—
Prior exposure	90	1071	10	—	—
Length of exposure	90	1070	10	—	—
Nonpotable wastewater					
Unadjusted	47	1044	5	3	1,5
Gender	47	1044	6	1	−1,3
Age	47	1043	6	2	0,4
Park activity	47	1043	4	5	3,7
Wet grass	47	1044	5	4	2,6
Sprayed	47	1042	4	4	2,6
Prior exposure	47	1043	6	4	2,6
Length of exposure	47	1041	4	6	4,8
Nonpotable run-off water					
Unadjusted	28	527	5	3	0,6
Gender	28	527	5	2	0,4
Age	28	491	5	3	0,6
Park activity	28	527	4	5	2,8
Wet grass	28	527	5	4	1,7
Sprayed	28	522	5	3	0,6
Prior exposure	28	526	7	3	0,6
Length of exposure	28	524	5	5	2,8

*Rate adjustment based on characteristics of total sample.
Gender: male (69), female (31).
Age: 0–15 years (19), 16–65 years (73), over 65 years (8).
Park activity: golf, softball, senior games (74), flag football, soccer (26).
Grass conditions: wet (40), dry (60).
Sprayed: yes (4), no (96).
Prior exposure: frequent (38), infrequent (62).
Length of exposure: less than 3 hours (63), more than 3 hours (37).

mean fecal coliform densities for subjects exposed to wastewater irrigation never exceeded 11.9/100 mL and 5/100 mL, respectively, while those for subjects exposed to runoff exceeded 12,000/100 mL and 3000/100 mL, respectively, in one study year. Similarly, total coliform counts per 100 mL averaged as high as 150 in one year for subjects in parks irrigated with wastewater compared to 60,000 for their runoff counterparts in one of the years. The densities to which subjects active in parks irrigated with potable water were exposed were a constant zero value.

Among subjects exposed to either wastewater or runoff irrigation, the percentage of symptom reports was found to be greater among those active in a park when the 500/100 mL fecal coliform density level was exceeded than when it was not exceeded (Table 2). Identical results were obtained for fecal streptococci at the 500/100 mL level and for total coliforms at the 3,000/100 mL level. Symptom reports were not found to differ significantly at any other bacterial density.

TABLE 2—Credible Symptom Reports by Fecal Coliform Counts/100 mL

	Under Count		Over Count			
Count/ 100 mL	% Reporting Symptom	Cases/ Total Subjects	% Reporting Symptom	Cases/ Total Subjects	Difference in Percent	95% Confidence Interval
200	4	37/922	6	28/439	2	0,4
500	4	37/953	7	28/427	3	1,5
3,000	4	46/1066	6	19/314	2	−1,5

Discussion

These results suggest the utility of fecal coliforms, fecal streptococci, and total coliforms at the identified density levels (500, 500, and 3000 per 100 mL, respectively) in establishing health standards for public park irrigation. Second, they suggest the advisability of park management practices that minimize the exposure of park patrons to wet grass conditions. And finally, they suggest that recycled water can be used for public park irrigation without undue hazard to health, provided that bacteria density levels are kept below those identified here.

ACKNOWLEDGMENTS

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New from nchs

First Births to Older Mothers, 1970–86

STEPHANIE J. VENTURA, AM

The number of first births to women in their thirties has quadrupled in the 16 years, 1970–86.[1] Throughout the 1970s, the increases were largest for women ages 30–34 but, since 1980, the increases in first births and first birth rates have been much larger for women ages 35–39.[1]

In 1986, 181,504 women ages 30–34 gave birth to their first child, more than four times the number (42,404) reported in 1970. Although the typical ages of childbearing continue to be 15–29 years, women in their early thirties accounted for 12 percent of all first births in 1986 compared with just 3 percent in 1970. The rise in first birth rates for these women has been nearly as impressive: 17.5 births per 1,000 women ages 30–34 in 1986 compared with 7.3 in 1970. Increases for women 35–39 were also substantial, although the numbers are considerably smaller. The rate for these women increased from 2.1 to 4.7 during this period (Table 1).

These findings are among those included in a recent report from the National Center for Health Statistics (NCHS) which also includes information on demographic patterns such as educational attainment and marital status of the mother and maternal and infant health characteristics.[1] The data are drawn from the live birth certificates of all States and the District of Columbia.

The increases during the early 1970s in first-time childbearing by women in their thirties occurred at the same time that there were steep reductions in higher order birth rates or, to put it another way, a decline in large families for older women. In the five years from 1970 to 1975, rates for fourth and higher order births for women aged 30 years and older fell by about one-half. As a consequence of these declines, which more than offset the slowly increasing first birth rates, the overall birth rate for women in their thirties dropped during the 1970s (Figure 1).

The reversal of these declines beginning in the mid-1970s is associated entirely with the sizable increases in first-birth rates, while higher order rates continued to decline but at a much slower pace. Higher order births comprised only 20 percent of all births to women aged 30 and older in 1986 compared with 55 percent in 1970.

The dramatic increases measured since 1970 in first-time childbearing by women aged 30 and older could have been anticipated because of several important demographic trends. First, there has been the widespread postponement of marriage by young people and the growing rate of dissolution of their marriages.[2,3] Thus, in 1986, 36 percent of women ages 25–29 years were not married (i.e., never married, widowed, or divorced) compared with 15 percent in 1970.[3]

A second important change has been the sharp decline and then leveling off in first birth rates for women in their twenties, particularly well-educated women, leaving large numbers and proportions of women still childless at age 30 and older. Between 1970 and 1975, first birth rates for women in their early twenties declined by nearly one-third; among women who were college graduates the rates fell by one-half. Some of this decline was made up for by small increases in birth rates for women ages 25–29. But the rates for women ages 25–29 stabilized by the late 1970s and, for college graduates, the rates have fallen since 1980. Because of these changes, large numbers and proportions of women have reached age 30 without having had any children. One-fourth of women ages 30–34 in 1986 were childless compared with 12 percent of comparably aged women in 1970.

A third important factor is the substantial growth in the number of women ages 25–39, from 18.4 million in 1970 to 30.9 million in 1986, a 68-percent increase.[4,5] These women, now ages 25–39 years, were all born in the baby boom years following World War II.

Fourth is the fact that national surveys show that most women intend to have at least one child; only about 10 percent of women in their early thirties expect to remain childless.[6]

Because of the interaction of these demographic patterns it was perhaps inevitable that the first birth rates for women in their thirties would surge upward in the late 1970s and 1980s. That is what has happened.

The educational attainment of women who became mothers at relatively older ages has dramatically increased; a larger proportion of these women than of women in any other age group has completed college. Since 1980, close to half of the women in their thirties having their first child were college graduates, compared with one-fourth of comparable women in 1970 (Figure 2). Birth rates specific for educational attainment provide further evidence that the postponement of childbearing and the making up of delayed first births have been primarily confined to well-educated women. For example, the first birth rate for college graduates ages 20–24 fell from 81 in 1970 to 29 in 1985, and for those ages 25–29 the rate fell from 80 to 58. In contrast, the first birth rate for college graduates ages 30–34 rose from 19 to 36 during these years and for women ages 35–39, the rate increased from 5 to 9 per 1,000 women (Table 2).

Greater educational attainment is associated with more timely receipt of prenatal care and a better outcome in terms of birthweight. For example, of first births to college graduates ages 30–34, only 5.8 percent were low birthweight (less than 2,500 grams or 5 pounds 8 ounces) compared with 7.5 percent of all first births to mothers in this age group.[1] In addition to seeking prenatal care very early in pregnancy, well-educated women are much more likely to be well-nourished and to gain

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Copies of the full report, Trends and Variations in First Births to Older Women, United States, 1970–86, Vital and Health Statistics Series 21, No. 47, are available from the U.S. Government Printing Office, Washington, DC 20402. Order stock no. 017-022-01081-0, price is $1.75.

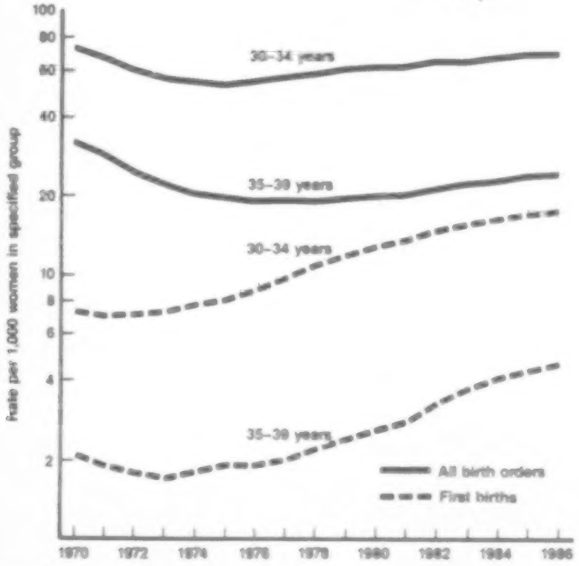
TABLE 1—Number of First Births and First Birth Rates by Age of Mother: United States 1970, 1980, 1985, 1986

	Age of Mother (years)						
Year	15–44[a]	15–19	20–24	25–29	30–34	35–39	40–44
				Number			
1970	1,430,680	498,388	652,530	212,102	4[illegible],404	11,704	2,442
1980	1,545,604	425,676	605,183	371,859	112,964	18,241	1,964
1985	1,554,788	359,272	552,974	418,658	170,686	39,447	3,804
1986	1,553,751	355,233	534,121	424,088	181,504	44,427	4,419
				Rate[b]			
1970	34.2	53.7	78.2	31.2	7.3	2.1	0.4
1980	29.5	41.4	57.3	38.2	12.8	2.6	0.3
1985	27.6	39.7	53.0	38.8	16.9	4.4	0.5
1986	27.2	39.1	52.7	38.8	17.5	4.7	0.6

[a] Numbers include births to women under age 15 years and ages 45–49 years which are not shown separately. Rates computed by relating first births regardless of age of mother, to women aged 15–44 years.

[b] First births per 1,000 women in specified group.

SOURCE: Reference 1, Tables 1 and 2.



SOURCE: Reference 1, tables 2 and 4

FIGURE 1—Birth Rates for Women Ages 30–34 and 35–39 Years, All Births and First Births, United States, 1970–86

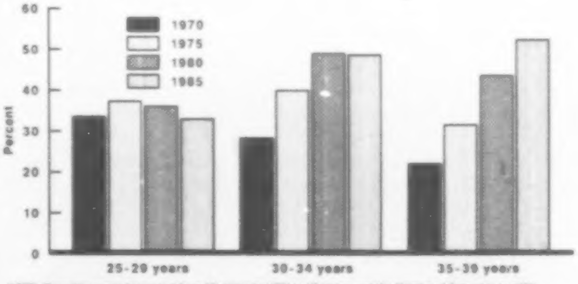


FIGURE 2—Percent of Women Having Their First Child Who Are College Graduates, by age

NOTE: Reporting areas increased from 38 states in 1970 to 47 states and the District of Columbia in 1985.

SOURCE: Reference 1, Table 7

TABLE 2—First Birth Rates for Women Who are College Graduates, by Age of Mother: United States 1970, 1975, 1980, 1985

	Age of Mother (years)			
Year	20–24	25–29	30–34	35–39
1970	81.1	80.3	19.1	5.0
1975	40.1	64.4	20.1	4.8
1980	39.7	65.8	29.8	6.6
1985	28.9	58.1	35.6	9.4

(Rates are first births per 1,000 women in specified age and educational attainment group)

Source: Reference 1, Table 8.

adequate weight during pregnancy as well as to curtail behaviors such as smoking which could affect pregnancy outcome.[7–9]

In recent years the number of babies born in the United States has been increasing because the number of women in the childbearing ages has risen. Birth rates by age of mother, however, have risen only for relatively older women.[10] The fact that a growing fraction of women giving birth for the first time are 30 years old or older may have implications for health care providers. For example, cesarean deliveries are more likely for older mothers.[11] Older mothers are at greater risk for certain complications,* and their babies are at higher risk of being of low birthweight and of having certain congenital anomalies.[10,12] Additionally, the much greater labor force participation rate of these older mothers may have implications in terms of child care needs. About two-thirds of first-time mothers ages 30–44 were in the labor force in June 1986 compared with half of women having their first child at ages under 25 years.[6] The data in this report can provide a basis for decisions on these medical, health, and child care issues.

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Errata

An error in the printing of Appendix Table 2 was recently found in the article by Sempos, *et al*, in the November 1988 Journal.[1] For 1979, the sum of the age-adjusted mortality rates of Ischemic Heart Disease and Hypertensive Heart Disease for black females should be 129.1. Also for 1979, the sum of the age-adjusted mortality rates of Ischemic Heart Disease, Hypertensive Heart Disease, and Cardiovascular Disease Unspecified for white males, white females, black males, and black females should be 244.9, 111.2, 254.2, and 151.0, respectively.

Also, the SOURCE should read: "Prepared by authors using data from the NCHS, Division of Vital Statistics." The editorial staff of the Journal regrets this error.

REFERENCE

1. Sempos C, Cooper R, Kovar MG, McMillen M: Divergence of the recent trends in coronary mortality for the four major race-sex groups in the United States. Am J Public Health 1988; 78(11):1422–1427.

* * *

In the paper by Lurie, *et al*,[1] May 1989 issue of this Journal, there is an error in the Table 2.

In column three, headed "Diff. (95% CI)," under the main heading of "Free", the % with natural impairment having an eye examination should read (3, 13) rather than (3, 3) as appeared in the published version. The authors regret that this error was not caught in the galleys.

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* * *

Concerning the letter to the Editor *Comments Received on Walker's 'Odd Man Out' Approach* published in the June 1989 issue of the *American Journal of Public Health* (Vol. 79, No. 6, p 781): The first sentence in the last paragraph should have read "It is not our aim . . ." rather than "It is our aim. . . ." Although it was felt that most readers can tell from the tone of the paragraph that the authors of the letter were not trying to detract from the Walker, *et al*, paper, they nonetheless requested that the Journal notify the readers of this error and express their regret over its occurrence.

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Effect of Disinfectants on Pseudomonads Colonized on the Interior Surface of PVC Pipes

ROGER L. ANDERSON, PHD, BETTY W. HOLLAND, JANICE K. CARR, BS,
WALTER W. BOND, MS, AND MARTIN S. FAVERO, PHD

Abstract: We investigated the effect of disinfectants on microbial contamination present on the interior surface of polyvinyl chloride (PVC) pipes filled with 600 ml of water contaminated with *Pseudomonas pickettii* and *P. aeruginosa*. After eight weeks, water was removed, and the test pipes exposed to various types of aqueous disinfectants. Disinfectant samples were removed, neutralized, and examined for recovery of microorganisms by membrane filtration. After seven-days exposure, disinfectant solutions were removed and pipes filled with sterile distilled water. Water was examined by membrane filtration at seven-day intervals to determine whether the organisms had survived in the pipes. Colonization of PVC surfaces were examined during each study phase by scanning electron-microscopy (SEM). *P. aeruginosa* was isolated directly from iodophor disinfectant, phenolic germicide, and iodophor antiseptic solutions. After addition of sterile water, *P. aeruginosa* was recovered from PVC pipes previously exposed to chlorine, phenolic, quaternary-ammonium, and iodophor disinfectants; *P. pickettii* was recovered from water in pipes treated with iodophor disinfectant, chlorine, and ethanol. The existence of glycocalyx-like cellular masses on the interior wall of PVC pipes most likely protected embedded organisms from the microbicidal action of some of the disinfectants tested and served as the reservoir for continuous contamination. (*Am J Public Health* 1990; 80:17–21.)

Introduction

The resistance of microorganisms to antimicrobial agents in general and chemical germicides in particular is controlled by a number of factors including the culture history and strain of the microorganism, the nature of the suspending menstruum as well as a variety of physical factors such as temperature, pH, and hardness. Further, microorganisms in their naturally occurring state have been shown to be significantly more resistant to chemical germicides than microorganisms that have been subcultured on artificial culture media.[1,2]

In the last several years, however, there have been reports of bacteria surviving in concentrations of chemical germicides that go significantly beyond the perceived limits of resistance. For example, recent epidemiologic and microbiologic investigations have documented intrinsic microbial contamination of iodophor antiseptic solutions.[3–6] Pseudobacteremia caused by *Pseudomonas cepacia* has been associated with the use of contaminated povidone-iodine, and *Pseudomonas aeruginosa* peritoneal infections has been attributed to the use of contaminated poloxamer-iodine. Presently, epidemiologic and laboratory investigations involving a contaminated povidone-iodine solution are in progress. Peritoneal infections in infants and false positive blood cultures from intensive-care unit patients with *P. cepacia* have been associated with the use of this intrinsically contaminated povidone-iodine antiseptic solution.[7] In addition, the prolonged survival of *P. aeruginosa* in iodophor solutions has been demonstrated after the iodophor was exposed to the inside surfaces of naturally contaminated PVC water distribution pipes.[8,9] When these organisms were isolated from iodine solutions and then tested, high numbers ($>10^6$ per ml) of *P. aeruginosa* were inactivated by iodophor solutions in minutes. More recently, studies have been completed which show recovery of *P. aeruginosa* from iodophors after unused or new PVC, stainless steel, copper, and glass pipes were artificially contaminated with water containing these organisms and subsequently exposed to iodine antiseptic solutions.[8]

The isolation and extended survival of microorganisms in these iodine-containing solutions may be due to extracellular glycocalyx-like material that microorganisms form and deposit on a variety of different types of surfaces.[10–12] Microbial colonization on the interior surfaces of PVC pipes and subsequent protection of organisms from the microbicidal action of antiseptic solutions may be due to the formation and adherence of extracellular matrixes to PVC surfaces. The end result is the production of thick masses of cells and extracellular material (biofilms). Once these masses are formed, antibacterial agents such as disinfectants or antiseptics must saturate the matrix before they can kill bacteria within adhesive biofilms.

Because PVC pipes are commonly used to distribute water in many pharmaceutical and manufacturing plants, in hospital dialysis units, and in various laboratory facilities, we initiated laboratory studies to determine: 1) the effects of various chemical germicidal disinfectants on *Pseudomonas* spp. colonizing the interior wall surfaces of new PVC pipe; 2) the observation and importance of thick cellular masses on PVC surfaces by SEM; and 3) if bacteria survive in and recolonize test pipes after exposure to chemical disinfectants. The purpose of these studies was to formulate a model system to determine the reason for the extraordinary ability of *P. cepacia* and *P. aeruginosa* to survive for extended periods of time after exposure to chemical germicides.

Methods

Preparation of PVC Test Pipes and Sample Remnants

Schedule 80 PVC pipe (4.1 cm in diameter) was used in these investigations. Test pipes consisted of two 30.4 cm PVC sections attached to a 90° PVC elbow. Straight pieces of PVC were separately attached to the elbow by pipe putty (Rectorseal, The Rectorseal Corp, Houston, TX), teflon tape (Teflon Thread Seal, DuPont, Wilmington, DE), organic solvents (Oatey PVC Cement, Oatey Co, Cleveland, OH), or dresser couplings (Flo-Control, Burbank, CA). Thus, four

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Editor's Note: See also related editorial p 13 this issue.

sections of PVC (distinguished by the type of attachment) were included in each test group. All PVC test pipes were thoroughly rinsed with tap water, dried, and separately packaged before ethylene oxide sterilization; rubber stoppers used to plug the pipe ends were steam sterilized.

Remnant samples for SEM were prepared by cutting small (1.5 by 1.5 cm) squares of PVC pipe. A small hole was bored at one end and a single 30.5 cm piece of monofilament fishing line tied to each PVC square. Remnant samples were packaged and sterilized with ethylene oxide before suspending them in PVC test pipes.

Microbiologic Conditioning of Test PVC Pipes and Remnant Samples

PVC test pipes (two test groups evaluated each week for four weeks) were filled with 600 ml of water containing a mixed, naturally occurring bacterial population which included only *P. aeruginosa* and *P. pickettii*. These organisms were originally present in naturally contaminated manufacturing plant water and served as the inocula for maintaining a source of contaminated water for these investigations. A reservoir of contaminated water was maintained by inoculating an 18-liter plastic carboy of sterile deionized water with 500 ml of contaminated plant water. Water in the carboy was held at ambient temperature (25°C), and contained a total microbial level of 10^4 colony-forming units (CFU) per ml at the beginning of this study. Initially, contaminated water present in the PVC test pipes contained 10^2 *P. aeruginosa* and 10^2 *P. pickettii* per ml of water.

Five remnant samples were suspended in rubber-plugged pipes filled with contaminated water. Pipes were then stored for eight weeks at 25°C to effect microbial colonization of the interior PVC surfaces. Microbial populations in the pipe waters were sampled weekly using membrane filtration procedures as described. Ten-fold water dilutions were made and passed through sterile sampling tubes attached to bacteriological membrane field monitors (0.45 micron; Millipore Corp, Bedford, MA). PVC remnants were removed at various time intervals for SEM observation of microbial colonization (i.e., formation of cellular and extracellular material accumulations).

Scanning Electron Microscopic Examination of PVC Remnants

For SEM studies, the PVC remnants were fixed in 1 percent glutaraldehyde in Sorenson's PO_4 buffer (pH 4.9) for 16 hours at room temperature, then rinsed three times in PO_4 buffer (pH 7.0). Postfixation was in osmium tetroxide (1 percent in PO_4 buffer, pH 7.0) for four hours at room temperature. The remnants were dehydrated in a graded ethanol series and then placed in a critical-point drying apparatus. Critical-point drying was carried out with CO_2 as the transitional fluid-drying gas. Specimens were loaded into the unit while immersed in 100 percent ethanol, and drying was accomplished by successive changes of liquid Co_2 while the apparatus was chilled by an ice/water bath. The CO_2 was removed by the temperature and pressure being raised to the critical point (32°C and 1100 psi) and slowly exhausted. Specimens were sputter-coated with gold palladium and examined in a Philips SEM 515.

Disinfection of Contaminated PVC Test Pipes

After eight weeks of incubation, contaminated water in the PVC test pipes was removed and cultured quantitatively using ten-fold dilutions and membrane filtration procedures as previously described. Two disinfectants were evaluated each week in two groups of PVC test pipes. Four contaminated PVC test sections (representing the four types of pipe attachment), or one PVC test group, were totally filled (*ca.* 600 ml) with each of the following disinfectants: 1) a 1:213 aqueous dilution (pH 2.9) of an iodophor detergent-germicide (Wescodyne, West Chemical Co, Lynbrook, NY); 2) a 1:128 aqueous dilution (pH 9.2) of a phenolic germicidal detergent (Vesphene II, Vestal Laboratories, St. Louis, MO); 3) a 1:256 aqueous dilution (pH 8.9) of a quaternary-ammonium germicidal detergent (Hi-Tor, Huntington Laboratories Inc, Huntington, IN); 4) a stock solution (pH 5.0) of an iodophor antiseptic (Prepodyne, West-Agro Chemical Co, Inc, Westwood, KS, manufactured for AMSCO Medical Products Division, Erie, PA); 5) 2 percent glutaraldehyde (Cidex, Surgikos, Inc, Arlington, TX) of pH 8.0; 6) a aqueous dilution (pH 7.5) of purified-grade sodium hypochlorite (Fisher Scientific Co, Fair Lawn, NJ) having 10–50 mg of free available chlorine per liter, as determined by the DPD colorimetric test method (Hack Company, Loveland, CO); 7) a 2 percent aqueous dilution (pH 3.4) of formaldehyde (Whiteworth Inc, Gardena, CA); and 8) 70 percent (vol/vol) aqueous ethyl alcohol (Warner-Graham Co, Cockeysville, MD) of pH 5.4. Two control PVC pipes were evaluated with each disinfectant test group: one noncontaminated PVC pipe with disinfectant and one water-contaminated PVC pipe with sterile distilled water.

Aliquots (10 to 50 ml) of the disinfectant were removed and sampled using membrane filtration procedures at 0, 2, 4, and 6 hours and 1, 2, 3, 5, and 7 days. After sampling, each pipe was refilled with an equal amount of the appropriate disinfectant. Filters were rinsed with 250 ml of brain heart infusion broth with 0.5 percent beef extract and the appropriate neutralizer. The filters were placed on blood agar plates. The following chemicals (i.e., in final percent concentration) were added to the enriched broth to neutralize any residual disinfectant: 0.5 percent sodium thiosulfate for iodophors and chlorine solutions; 0.5 percent Tween 80 for phenolics; 0.7 percent lecithin (soybean) for quaternary ammonium compounds; and 0.25 percent sodium bisulfite for aldehydes. Blood agar plates with filters were incubated for 24–48 hours at 35–37°C. The microbial count was recorded per aliquot sampled and subsequently expressed as CFU per ml. To insure that *P. aeruginosa* and *P. pickettii* were the only organisms in the mixed microbial population that was used to condition the PVC pipes, two representative colonies of each morphologic type were picked from the filters and inoculated to triple sugar iron agar slants. Organisms showing nonfermentative reactions on these slants were identified by standard biochemical tests.[13]

Recolonization of PVC Pipes Previously Exposed to Disinfectants

After the contaminated pipes were exposed to disinfectants for seven days, the pipes were emptied and each was refilled with 600 ml of sterile distilled water. Pipes were resealed with fresh sterile rubber stoppers. Water was removed at seven-day intervals and various ten-fold dilutions were examined by membrane filtration procedures as previously described to determine reestablishment of microbial contamination in PVC test pipes previously exposed to chemical disinfectants. After dilutions were filtered, the membrane filters were placed directly on blood agar plates (no neutralizer rinse) and incubated.

Results

Microbiologic Conditioning of PVC Test Pipes

Over the eight-week conditioning period, water counts for *P. pickettii* (biovar Va-1) were higher (10^5 to 10^6 CFU per ml) than counts observed for *P. aeruginosa* (10^3 to 10^5 CFU per ml). The microbial count for each organism increased each week during the eight-week sampling time. Total viable counts were approximately ten-fold higher in PVC test pipes that used pipe putty (a pipe-threading compound) to attach individual straight pieces to the 90° PVC elbow than pipes that did not have putty.

SEM Examination of PVC Remnants Exposed to Contaminated Water

The presence of bacterial colonization (i.e., formation of cells embedded in extracellular material) on the interior surface of PVC sample remnants was observed by SEM after remnant exposure to contaminated water (Figure 1). Microorganisms were not observed on control PVC remnants (samples not exposed or conditioned with contaminated water).

Recovery of Organisms from Test Disinfectants

P. aeruginosa was recovered directly from an iodophor disinfectant and a phenolic germicidal detergent in 48 hours and from an iodophor antiseptic solution in five days (Table 1). *P. pickettii* was not isolated from any of the test chemicals through seven days of sampling. *P. aeruginosa* and *P. pickettii* organisms were recovered from each of the 32 sterile water controls (i.e., four controls for each test disinfectant). The method of attaching PVC sections to the 90° PVC elbow did not correlate with recovery of organisms from test disinfectants.

Reestablishment of Organisms in Sterile Water

P. aeruginosa was recovered from water placed in PVC pipes previously exposed for seven days to use dilutions (i.e., diluted product as recommended by the manufacturer) of either an iodophor disinfectant, a phenolic germicidal detergent, a quaternary-ammonium compound, an iodophor antiseptic, or a dilution of sodium hypochlorite; *P. pickettii* was recovered from the pipes treated with an iodophor disinfectant, a chlorine solution, or 70 percent ethanol (Table 2). There was no evidence of microbial contamination during 42 days of sampling from contaminated PVC pipes exposed to 2 percent glutaraldehyde or 2 percent formaldehyde and refilled with sterile water. In some instances, organisms were found in water present in the PVC test pipes but were not recovered from the corresponding disinfectant initially used to disinfect that group of pipes.

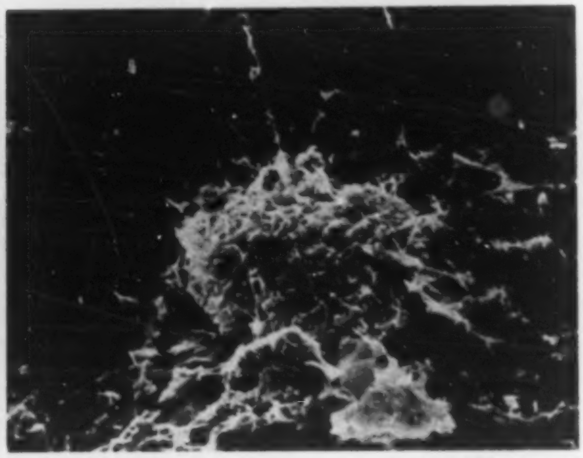

FIGURE 1—Scanning Electron Micrograph of the Interior Surface of a PVC Remnant Sample after Eight Weeks of Exposure to Contaminated Water, × 3,800

Discussion

Previous investigations have demonstrated the prolonged survival of *P. aeruginosa* in naturally contaminated PVC water distribution pipe segments after they were exposed to iodophor antiseptic solutions.[8,9] In one study, *P. aeruginosa* was isolated from a poloxamer-iodine solution 48 hours after membrane-filter sterilized poloxamer-iodine was added to the interior surface of naturally-contaminated PVC water pipe. Continuous exposure of poloxamer-iodine to this pipe resulted in a level of 10^4 CFU per ml of *P. aeruginosa* at nine days. This contaminated solution of poloxamer-iodine was removed from the pipe, stored at 25°C, and served as the reservoir for prolonged microbial survival studies. Using this contaminated stock solution of poloxamer-iodine, we found that *P. aeruginosa* (10^4 CFU per ml) survived for as long as 98 days in stock solution after removal from the PVC pipe.[8] Based on this investigational approach, a suitable laboratory model for induction of survival of these organisms and for studying the mechanism of *Pseudomonas* resistance to iodophors was developed.

In addition, studies have been done in which unused pipes of different composition were artificially contaminated by exposure to contaminated water containing *P. aeruginosa*.[8] After seven days, the water was removed from each pipe, and all were filled with membrane-filtered poloxamer-iodine. *P. aeruginosa* was recovered from poloxamer-iodine after exposure to artificially-contaminated PVC, stainless steel, copper, and glass pipes, but not galvanized steel pipe.

Thus, the newly developed model of using both naturally and artificially contaminated PVC pipes proved useful in demonstrating iodophor-resistance of *P. aeruginosa*. To date, the addition of iodophor antiseptic solutions to contaminated PVC pipe is the only laboratory method now available to show recovery and survival of *P. aeruginosa* from iodophor antiseptics.

We report experiments designed to determine: 1) the effect of various germicides on *Pseudomonas* spp. colonizing PVC pipe; 2) the existence of glycocalyx-like structures as a potential survival mechanism for these organisms; and 3) the reestablishment of *Pseudomonas* spp. in sterile distilled water added to contaminated and disinfectant-treated PVC pipes. *P. aeruginosa* was recovered from three test germicides during a seven-day exposure period: an iodophor disinfectant and phenolic germicide in 48 hours and an iodophor antiseptic in five days. *P. aeruginosa* was not recovered from either a quaternary-ammonium germicidal detergent, 2 percent glutaraldehyde, 10–50 ppm chlorine, 2 percent formaldehyde, or 70 percent ethyl alcohol; *P. pickettii* was not isolated from any of the eight disinfectants tested. The variable microbicidal capabilities of the germicides tested could be due to the type and amount of bacterial colonization present in the PVC pipes, the presence of morphologically abnormal bacterial cytoplasm in cells, or the specific mode and site of action of test disinfectants on bacterial cells (such as microorganisms embedded in a matrix of extracellular material).

Recent review articles have discussed contaminated antiseptics and disinfectants and the various types of orga-

TABLE 1—Recovery of Organisms from Chemical Disinfectants Exposed to Contaminated PVC Test Pipes

Test Pipe Group and Disinfectant Used	Recovery of Organisms per Aliquot of Chemical Sampled:		Recovery of Pseudomonads from Sterile Water Controls:	
	P. aeruginosa	*P. pickettii*	*P. aeruginosa*	*P. pickettii*
A[a]-Iodophor Disinfectant	48h, 2/50 ml	–[b]	+[c]	+
B-Phenolic Germicide	48h, 4/50 ml	–	+	+
C-Quaternary Ammonium Comp.	–	–	+	+
D-Iodophor Antiseptic	5d, 39/10 ml	–	+	+
E-2% Glutaraldehyde	–	–	+	+
F-Chlorine (10-50 ppm)	–	–	+	+
G-2% Formaldehyde	–	–	+	+
H-70% Ethanol	–	–	+	+

[a] Each group consisted of four test PVC pipes.
[b] No recovery after seven days of chemical exposure.
[c] Recovery of organisms from sterile water controls of each disinfectant test group (a total of 32 positive controls).

TABLE 2—Reestablishment of Organisms in Sterile Distilled Water after Exposure of Contaminated PVC Pipes to Disinfectants

Test Pipe Group and Disinfectant Used	Recovery of Organisms from Water after Pipes Were Exposed to Chemicals	
	P. aeruginosa	*P. pickettii*
A[a]-Iodophor Disinfectant	7d, 3.1×10^5/ml	7d, 7.5×10^4/ml
B-Phenolic Germicide	7d, 7.5×10^5/ml	–[b]
C-Quaternary Ammonium Comp.	7d, 25/50 ml	–
D-Iodophor Antiseptic	7d, 2.8×10^2/ml	–
E-2% Glutaraldehyde	–	–
F-Chlorine (10–50ppm)	7d, 2.1×10^3/ml	7d, 3.5×10^5/ml
G-2% Formaldehyde	–	–
H-70% Ethanol	–	21d, 1.4×10^6/ml

[a] Each group consisted of four test PVC pipes.
[b] No recovery after 42 days of water sampling.

nisms associated with these contaminated chemical solutions[14] as well as the resistance of bacteria to antiseptics and disinfectants.[15] A more definitive answer as to why *Pseudomonas* may survive or even grow in germicidal solutions will most likely not be available until laboratory studies are completed which clarify the mechanism of germicidal activity and determine the biologic and ecologic characteristics of *Pseudomonas* spp. that permit their tolerance to or utilization of various chemical compounds.

Massive concentrations of bacterial cells were observed by SEM on the surface of PVC remnant samples after incubation in contaminated water. These cellular masses (or glycocalyx-like structures) appeared to develop in greater numbers as the length of exposure time to contaminated water increased. As suggested in this investigation, the existence of glycocalyx-like structures and the shedding of these structures from interior PVC wall surfaces most likely served to shield *P. aeruginosa* from the microbicidal action of iodophor and phenolic germicides (Table 1).

The protective effect of the bacterial glycocalyx has been suggested by several investigators.[8–12,16] But it should be noted that our preparative techniques for SEM observations were not sufficient to preserve the entire glycocalyx matrix as described by Coseterton, *et al.*[10] Although we observed a considerable degree of microbial adhesion accompanied by notable accumulations of extracellular material on the PVC surfaces, our visual observations are likely an underestimation of the actual amount of physical mass. The prolonged survival of *Serratia marcescens* in 2 percent chlorhexidine was most likely due to these microorganisms being embedded in a thick matrix which adhered to the walls of storage jugs containing this antiseptic solution.[17] Viable *S. marcescens* was recovered from 2 percent chlorhexidine during a storage period of 27 months. The recovery of *P. aeruginosa* from unopened bottles of poloxamer-iodine could be the result of contaminated water and PVC product distribution pipes from a manufacturing plant.[16] SEM micrographs of these pipes demonstrated large concentrations of rod-shaped and coccobacillary cells embedded in interior deposits of the pipe. Also, large numbers of *P. aeruginosa* (10^4 CFU per ml) survived for as long as 98 days in a stock solution of poloxamer-iodine after exposure and removal from naturally contaminated PVC water distribution pipe.[8]

In the present study, microorganisms were observed to survive and reestablish in sterile water placed in PVC pipes previously exposed to chemical germicides. Although *Pseudomonas* spp. were not recovered from some of the test disinfectants, they were isolated from water that was placed in the corresponding chemically-treated test pipes. This reestablishment of microbial contamination suggests a continuous reservoir of organisms adhering to and shedding from the interior PVC wall surface. Organisms showing survival in five PVC test pipes were *P. aeruginosa* (after exposure to chlorine or a quaternary ammonium compound) and *P. pickettii* (after exposure to a iodophor disinfectant, chlorine, or 70 percent ethanol). This is strong evidence that these organisms survived within the colonized PVC pipes after exposure for seven days to a variety of chemical treatments. The physical thickness of cellular and extracellular material that form on PVC pipe surfaces could protect organisms from the action of germicidal chemicals and serve as a continuous reservoir for microbial contamination in test pipes and of water flowing through pipes.

Other investigators have hypothesized the importance of cell masses such as the glycocalyx as a potential reservoir of microbial contamination.[18,19] Several recent investigations have implicated *P. aeruginosa* as the etiologic agent of whirlpool-associated infections (i.e., folliculitis). These organisms are capable of growing in water and colonizing various whirlpool surfaces including water pipes. Such cellular and extracellular material accumulations adhering to whirlpool surfaces could protect bacteria from the action of chlorine and provide a constant microbial reservoir resulting in the contamination of water after a chemical treatment.

The public health significance of this investigation and its findings are widespread. PVC pipes are commonly used in laboratory facilities, pharmaceutical and manufacturing plants, hospitals, therapeutic and public whirlpool-type hot tubs, and homes to distribute water for various usages. In 1984, during an investigation of a manufacturing plant which produced iodophor antiseptics, *P. aeruginosa* contamination was recovered throughout the PVC water distribution system.[16] Subsequently it was found that intrinsic contamination of a poloxamer-iodine antiseptic solution resulted when the formulated iodophor was allowed to stand in contaminated PVC pipes (pipes between the mixing tank and storage tank and between the storage tank and the bottling area) prior to bottling. As remedial measures, the company removed all PVC pipes, installed new stainless pipe, and initiated a hot water sanitizing procedure (60°C water for one hour) of all process water and product distribution pipes. In addition, better and more frequent microbiological quality control procedures on municipal and process water were instituted.

These types of laboratory studies are important in identifying possible mechanisms by which microorganisms could survive in disinfectant solutions for extended periods of time and to determine how bacterial organisms could continuously contaminate the various parts within a water distribution system. The best way to maintain a water distribution system in a safe condition is to limit the number of microorganisms by proper and scheduled maintenance and to routinely sanitize pipes or tubing that transport water. The laboratory findings presented here add to development of effective strategies for disinfecting water or product distribution lines and holding tanks in manufacturing plants and in laboratory, medical, and pharmaceutical facilities where microbial contamination by pseudomonads or other naturally-occurring gram-negative water bacteria is a problem.

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Public Health Workforce Consortium Recommended

Development of a Public Health Workforce Consortium with public and private representatives—to study health personnel implications of strategic public health issues and improve data collection efforts—has been recommended by a work group established by the Health Resources and Services Administration (HRSA).

The HRSA-supported group recommended the study of selected issues such as those delineated by the Year 2000 Health Objectives for the Nation. It also recommended an incremental approach to the eventual creation of an analytical system capable of assessing the public health work force.

Information on the Public Health Workforce Consortium project and its recommendations is available from B. Jerald McClendon, Division of Associated and Dental Health Professions, Bureau of Health Professions, Room 8C-09, DHHS/PHS, 5600 Fishers Lane, Rockville, MD 20857. Telephone: (301) 443-6757.

First Nursing Home Admissions: Time Spent at Home and in Institutions after Discharge

Mary Ann Lewis, RN, DrPH, Barbara Leake, PhD, Margaret Leal-Sotelo, MSW, and Virginia Clark, PhD

Abstract: We tracked 96 percent of a sample of 1,942 nursing home patients admitted to a nursing home for the first time in 1982–83. Patients discharged alive from the nursing home were followed for two years or until death. The relative time spent at home, in hospitals, and in skilled nursing facilities is reported.

Of the 705 patients discharged from their initial nursing home admission to homes in the community, about 50 percent made only one transfer and only 15 percent made four or more transfers. Of the 509 discharged to a hospital, 26 percent died there and 37 percent of the 374 survivors made four or more moves in the next two years. In all, 1,332 patients were discharged alive and they spent almost two-thirds of the subsequent two years, or their remaining lifetimes, in the community. Of those who transferred only once, over two-thirds of their follow-up time was spent in their own homes. Policies concerned with long-term care should use some type of actuarial data base to successfully plan and implement long-term care insurance. (*Am J Public Health* 1990; 80:22–24.)

Introduction

As the United States population continues to age, there is growing concern over the costs of long-term care and the importance of developing mechanisms to finance such care.[1,2] Projections of the costs of an insurance program should be based on data that includes patients' use of all types of services and facilities in the health care system, their survival, and the resources required to provide services to maintain them at home. Despite the size of the long-term care segment of the health care system, and the need for its services, relatively little is known about patients' continued use of services, especially non-institutional services, until death.[3]

Our previous longitudinal studies of defined nursing home populations examined the immediate discharge outcomes of nursing home care in three cohorts of elderly patients (1980, 1982–83, 1984).[4–6] Over time there were no changes in casemix or immediate discharge outcomes of persons admitted to nursing homes for the first time (first admissions).[5] However, a comparison of first admissions and readmissions showed that readmissions were sicker and had worse short-term outcomes.[6] A comparison of readmissions in 1980 and 1984 showed 1984 readmissions were more debilitated than those in 1980 and less likely to return home.[7] Further, readmissions who had been admitted from hospitals in 1984 were almost twice as likely to be rehospitalized within 30 days compared to all readmissions in 1980.

Using a small cohort of patients in 1980, we did a preliminary examination of the number of post-discharge transfers for two years after the initial discharge from the nursing home.[8] In the present study, we provide a detailed report on the continued use of services including time spent at "home" and in institutions, over a two-year period using a 1982–83 discharge cohort of first admissions to nursing homes.

Methods

A cohort of 1,942 patients discharged in 1982–83 from 45 non-hospital based skilled nursing facilities (SNFs) in the urban areas of San Bernardino/Riverside/Ontario in Southern California comprised the total sample. A detailed account of the data collection methods for this first nursing home admission sample has been reported.[5]

The initial discharge outcomes were classified as:

- returned to home in the community,
- transferred to another nursing home,
- died in the nursing home,
- transferred to the hospital and died there, and
- transferred to the hospital and survived (i.e., subsequently discharged alive from the hospital to a nursing home or to a home in the community).

Patients discharged alive were followed for a period of two years or until death following their initial discharge from a nursing home. The location and associated follow-up information for patients discharged alive were obtained from hospital and nursing home medical records, correspondence with medical record librarians, and telephone interviews with family and friends. This patient relocation information was obtained for two years after the initial nursing home discharge by recording each patient's change of setting; length of stay in that setting, and whether the patient was discharged alive or dead from that setting. Time at home in the community was separated into days spent in the patient's "own home," which included their personal home or that of a relative or friend, days in a retirement home, and days spent in a board and care facility.

Length of life was calculated from the time of discharge from the initial nursing home until death or the end of the two-year follow-up period. Deaths were verified in the county vital statistics records.

This paper describes the associations between number of transfers (0, 1, 2, 3, and 4, or more) and patients' immediate discharge outcomes. It also provides detailed documentation on patients' length of stay in the initial SNF, all subsequent SNFs, hospitals, and homes, including their own (or that of friends or relatives), as well as retirement and board and care homes, according to the number of transfers. Because distributions of survival are skewed, both medians and means are reported.

Results

We successfully tracked 96 percent (N = 1865) of the

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total sample of patients (N = 1,942) for the two-year follow-up. Those patients lost to follow-up (N = 77) were excluded from the analysis. At two years, a review of county death records showed that none of the 77 patients lost to follow-up had died in the study area.

Table 1 shows the association between the initial nursing home discharge outcome and the number of subsequent transfers (0–4 or more) for two years. Over one-fourth (N = 533) of the total patients tracked (N = 1,865) died in the nursing home. Almost 30 percent (N = 550) made only one transfer; of these, 11 percent moved to other SNFs, and 64 percent returned to homes in the community. The remainder (25 percent) went to hospitals where they died.

About 18 percent (N = 342) of all patients made two transfers after discharge from the SNF. Of these, 50 percent were discharged to hospitals and survived; 44 percent initially returned to homes in the community; the rest moved to another SNF. Ten percent (N = 187) of all patients made three transfers; of these 16 percent initially went to other SNFs; 35 percent were discharged to hospitals and lived; and almost half returned to homes in the community. The remainder of the patients (N = 253 or about 14 percent) made four or more transfers; only 3 percent of this group were discharged to other SNFs; 55 percent went to hospitals and survived; and 42 percent returned initially to homes in the community.

Put another way, of the 705 patients discharged from their initial nursing home admission to homes in the community, 354 (about 50 percent) made only one transfer and only 15 percent (106/705) made four or more transfers. Of the 509 discharged to a hospital, 26 percent died there; and of the 374 survivors, 37 percent made four or more moves in the next two years.

Table 2 summarizes the survival time of each group making at least one transfer, the proportion alive at two years, and the relative distribution of days of life spent in various locations, i.e., SNFs, hospitals and homes in the community after discharge from the initial nursing home.* An overall summary is also given for all patients discharged alive. The median survival time after discharge for the entire cohort of patients was 157 days. After eliminating those individuals who died in the nursing home, the median survival time was 617 days. Patients discharged alive from the nursing home spent 65 percent of their remaining time during the two-year follow-up in the community. Following discharge from the nursing home, these individuals spent approximately 33 percent of their time in SNFs, 3 percent in hospitals, 11 percent in board/care homes, 3 percent in retirement homes, and 51 percent in their "own" homes.

TABLE 2—Distribution of Time Spent in Various Locations after SNF Discharge and Survival for Two Years According to Number of Transfers

	Number of Transfers				
	1	2	3	≥4	All
No. of Patients	550	342	187	253	1,332
Median LOS in Initial SNF	33	28	36	29	32
% Alive at Two Years:	53	33	49	53	47.2
Median Survival in Days:*	—	293	691	—	617
(Average)**	444	360	483	538	448
% Days Spent In:					
SNF	13.0	49.5	38.1	50.1	32.6
Hospital	0.5	3.1	4.6	5.5	2.7
Board/Care	15.0	7.2	7.6	7.0	10.4
Retirement Home	4.0	2.5	3.0	1.0	2.9
Own Home	67.5	37.7	46.7	36.4	51.4
Any Home***	86.5	47.4	57.3	44.4	64.7

*—Unable to compute median due to censored data related to long-term survivors.

**Censored data causes average survival to be biased because information is lacking on long-term survivors.

***"Any Home" includes all types of homes, i.e., board and care, retirement patient's own home, and the home of a friend or relative.

The median survival in days was 293 for those who transferred twice,and 691 days for those who made three transfers.[9] Because of censored data related to missing information on long-term survivors, we were unable to compute the median survival for individuals who made one and four or more transfers.

As indicated in Table 2, only those making either two or four or more transfers spent the majority of their lives during the two-year follow-up period in institutions. Of those making only one transfer, 53 percent were alive at two years, and as a group they spent over 86 percent of their follow-up days in the community, with over 67 percent of those community days being spent in their "own" homes. For those making three transfers, over half of their post SNF discharge days of life during follow-up were also spent in the community, and the vast majority of that time in the community was spent in the patients' "own" homes. Of the 1,332 patients in Table 2 who spent variable proportions of their remaining days of life at home, 56 percent were disoriented, 46 percent were incontinent of stool and/or urine and 76 percent were dependent in activities of daily living.

TABLE 1—Association between Initial Nursing Home Discharge Outcome and Number of Post Discharge Transfers for Two Years

				Transferred from SNF to:			
					Hospitals		
Number of Transfers	Number of Patients	Percentage of Patients	Died in SNF	SNF	Died	Lived	Home*
0	533	28.6	533	0	0	0	0
1	550	29.5	0	61	135	0	354
2	342	18.3	0	19	0	171	152
3	187	10.0	0	29	0	65	93
≥4	253	13.6	0	9	0	138	106
Totals	1,865	100%	533	118	135	374	705

*Home includes all types of homes, i.e., board and care, retirement, patient's own home, and the home of a friend or relative.

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*Note that "any home" in the Table includes board/care, retirement and "own" homes.

Discussion

This study documents the subsequent number of transfers and resource utilization of a discharge cohort of 1,865 patients admitted to skilled nursing facilities for the first time. Almost 30 percent of the sample died in the nursing home, but among those who survived, an unexpected and interesting finding was the length of time many of these patients spent in homes (usually their "own") in the community after being discharged.

The transfer of the frail elderly with variable needs for continued care from one setting to another is a function of several factors. Clearly health status or requirements for care affect placement, especially for those with the greatest need for 24-hour care. The supply of beds, reimbursement policies and patients' economic resources interact to limit length of stay in different facilities. Finally, the preferences of patients and families may well be the determinant of placement at home or in a long-term care facility for those with the personal and or economic resources to exercise this option. This study clearly demonstrates the magnitude of "at home" care utilized by those who survive their first nursing home admission. There are no reported studies, to date, of the relative importance of the variables listed above in determining the mobility of this population, and the non-institutional costs of providing care for them.

Thirty percent of these patients made just one transfer after discharge, and then spent over two-thirds of the succeeding two years, or their remaining lifetime in their own home or that of a friend or relative. There is no reported information related to the family and community resources required for this amount of home care. Over one-half of the individuals making one transfer survived the two-year follow-up period. This represents a considerable amount of time at home. Given the burden of care requirements these patients present, one can only speculate about the degree to which their nursing care requirements were met, the impact of their need for care on family resources, and how (already mentioned this in terms of degree) the challenge was partitioned among family, friends, and other home care providers. Clearly, the need for a more detailed study of patients' continuing care needs (and how these were met) is apparent. Whether or not the families of these individuals managed independently to care for their loved ones or used community nursing agencies is important information for those concerned with policy. Another economically relevant issue that deserves consideration is lost work time for family caregivers.

Information on the specific care requirements for the significant portion of the persons in the sample who were living at home two years after discharge is particularly essential. Such data (while not collected in this study) can best be derived from longitudinal tracking of patients, a research design not commonly used in nursing home studies.[10]

As indicated in previous publications, the area for this study represents the 37 standard statistical metropolitan (SMSA) areas throughout the country with the exception that it has a lower proportion of Blacks and individuals below the poverty level. It may or may not represent the population in the eight urban SMSAs, e.g., New York City, Los Angeles, San Francisco, etc.

This data set does provide a fairly detailed picture of the natural history of continuing care for a modest sample of patients for two years after discharge from nursing homes. Estimates of costs may be derived from the number of days spent in various institutional settings (hospitals, skilled nursing facilities, board and care homes). Advocates of long-term care insurance will require some type of actuarial data base of individual patients' utilization of facilities and resources to successfully plan and implement long-term care insurance. However, what is conspicuously missing from this report is information on the services provided to these individuals and their social and economic costs during those intervals spent at home. More information of this type is essential to planning for the care of the rapidly growing numbers of frail elderly, and assisting families to cope with the burdens of their care.

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Cocaine and the Risk of Low Birth Weight

Diana B. Petitti, MD, MPH, and Charlotte Coleman, BA

Abstract: In a population-based case-control study, we assessed the relation between cocaine use and the risk of low birth weight in Alameda County, California, excluding women who used heroin and/or methadone during pregnancy. After adjustment for confounders, the relative risk of low birth weight in users of cocaine throughout pregnancy was 4.0 (95% CI = 1.8, 8.9). Cocaine use was estimated to account for 10 percent of cases of low birth weight in Black women in Alameda County. (*Am J Public Health* 1990; 80:25–28.)

Introduction

Cocaine use affects pregnancy outcome adversely in animals.[1,2] In humans, it has been associated with an increased risk of abruptio placenta[3,4] and of intrauterine growth retardation and pre-term delivery.[5–11] In a population-based case-control study of low birth weight in Alameda County, California, we estimated the relative risk of low birth weight in relation to use of cocaine and the contribution of cocaine use to the problem of low birth weight in this urban population.

Methods

Subjects were identified using information from Alameda County, California birth certificates. Eligible were singleton infants without congenital anomalies whose mother and father were both Black or both White, non-Hispanic, born in Alameda County between January 1, 1987 and December 31, 1987 to residents of Alameda County. Cases weighed 500 through 2499 grams. Controls were chosen at random from among infants with weights of 3000 or more grams, the number of controls in each ethnic group being equal to the number of cases.

We identified 500 Black cases and 500 Black controls and 268 White, non-Hispanic cases and 268 White, non-Hispanic controls for the study, but nine records were subsequently judged ineligible. After obtaining permission from the California Vital Statistics Advisory Committee to use information from the confidential section of the birth certificate and the approval of our study protocol and interview instrument by two local committees for protection of human subjects, we used information on the birth certificate to locate women, who were given a detailed explanation of the study's aims, of the risks and benefits of participation, along with assurances of the confidential nature of the information provided. Consenting women were interviewed in their homes by trained interviewers. Ultimately, 766 of 994 eligible Black mothers (77.1 percent) and 462 of 533 eligible White mothers (86.7 percent) were interviewed.

Women were separately queried about use of cocaine and of "crack" or freebase cocaine, which is the smoking of pure crystalline cocaine.[12] Information about other potential confounders was also collected by interview.

Cases of low birth weight were sub-classified into three groups: ≤1500 grams (very low birth weight), 1501–2500 grams and <37 menstrual weeks of gestation (low birth weight, pre-term), and 1501–2500 grams and ≥37 menstrual weeks of gestation (low birth weight, term, intrauterine growth retarded) according to suggestion of Yerushalmy.[13]

The relative risk of low birth weight was estimated using the exposure odds ratio. The chi-square statistic was used for polychotomous variable. Ninety-five percent confidence intervals were estimated using the method of Cornfield,[14] except when these failed to converge, in which case the method of Kleinbaum and Kupper was used.[15] Logistic regression as implemented on SAS[16] was used to estimate relative risk controlling for confounders.

The population attributable risk percent was estimated using methods described by Walter.[17] The population prevalence of cocaine use in Whites and Blacks was estimated as the sum of the number of cases who used cocaine and the number of controls who used cocaine multiplied by the relevant sampling ratio (26.32 for Whites and 7.75 for Blacks) divided by the estimated total number of White or Black deliveries in Alameda County.

Results

Seven Black women and four White women who used heroin and/or methadone during pregnancy, and four Black women with incomplete information on cocaine use, were excluded from all analyses. After these exclusions, both the use of cocaine in forms other than crack throughout pregnancy and the use of crack throughout pregnancy were associated with large increases in the risk of low birth weight (Table 1).

No White women admitted to use of crack or cocaine in other forms throughout pregnancy, and multivariate analysis was confined to Blacks. Because of the small number of controls who admitted to use of cocaine or crack throughout pregnancy, we combined the categories for cocaine and crack use throughout pregnancy. After adjustment for confounders, the increased relative risk of low birth weight in women who used cocaine and/or crack throughout pregnancy persisted but was attenuated (Table 2).

Low pre-pregnancy weight and low weight gain may be caused at least in part by use of cocaine, and a history of having a prior low birth weight infant may be related to use of cocaine in the prior pregnancy. Thus, a multivariate statistical model that includes these factors may "over

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TABLE 1—Estimated Relative Risk of Low Birth Weight by Cocaine Use in Blacks and Whites

	Blacks			Whites		
Variables	N Cases	N Controls	RR (95% CI)	N Cases	N Controls	RR (95% CI)
Self-Reported Cocaine Use during Pregnancy						
Did not use at all	297	375	1.0[a]	218	234	1.0[a]
Used cocaine and/or crack, not throughout pregnancy	14	5	3.5 (1.2, 11.4)	2	4	0.5 (0.1, 3.0)
Used crack with or without cocaine in other forms throughout pregnancy	32	4	10.1 (3.5, 28.9)	0	0	—[b]
Used cocaine but not crack throughout pregnancy	23	5	5.8 (2.1, 17.6)	0	0	—[b]

[a] referent

[b] cannot be estimated

TABLE 2—Results of Three Multiple Logistic Regression Analyses Examining Cocaine/Crack Use and the Relative Risk of Low Birth Weight Controlling for Various Other Variables[a]

Variables	Full Model RR (95% CI)	Excluding Weight Gain <15 Pounds RR (95% CI)	Excluding Weight Gain <15 Pounds, Low Pre-Pregnancy Weight RR (95% CI)
Self-Reported Cocaine/Crack Use during Pregnancy			
Not at all	1.0[b]	1.0[b]	1.0[b]
Used cocaine and/or crack but not throughout pregnancy	1.8 (0.5, 6.4)	2.4 (0.7, 7.9)	2.3 (0.7, 7.6)
Used cocaine and/or crack throughout pregnancy	4.0 (1.8, 8.9)	4.4 (2.0, 9.7)	4.4 (2.0, 9.6)
Cigarette Smoking during Pregnancy			
Not at all	1.0[b]	1.0[b]	1.0[b]
Used but quit	1.3 (0.6, 2.6)	1.3 (0.7, 2.5)	1.3 (0.7, 2.5)
Used throughout	3.1 (2.1, 4.5)	3.2 (2.2, 4.7)	3.3 (2.2, 4.7)
Maternal Age (years)			
<17	1.7 (0.7, 3.8)	1.8 (0.8, 3.9)	1.8 (0.8, 3.8)
17–34	1.0[b]	1.0[b]	1.0[b]
35+	1.9 (1.0, 3.5)	2.0 (1.1, 3.7)	1.9 (1.0, 3.5)
Prior Live Births			
0	2.3 (1.5, 3.3)	2.1 (1.4, 3.1)	2.2 (1.5, 3.2)
1–2	1.0[b]	1.0[b]	1.0[b]
3+	1.4 (0.8, 2.5)	1.2 (0.7, 2.2)	1.3 (0.7, 2.3)
Alcohol Use during Pregnancy			
<3 drinks/day	1.0[b]	1.0[b]	1.0[b]
3+ drinks/day	1.7 (0.7, 4.1)	1.5 (0.6, 3.6)	1.4 (0.6, 3.5)
Socioeconomic Status Index[c]			
High	1.0[b]	1.0[b]	1.0[b]
Middle	1.1 (0.7, 1.7)	1.2 (0.8, 1.8)	1.2 (0.8, 1.9)
Low	1.5 (0.9, 2.3)	1.7 (1.1, 2.7)	1.7 (1.1, 2.7)
Prior Low Birth Weight Baby			
No	1.0[b]	1.0[b]	1.0[b]
Yes	6.9 (3.7, 13.2)	7.5 (4.0, 14.0)	7.6 (4.1, 14.1)
Low Pre-pregnancy Weight[d]			
No	1.0[b]	1.0[b]	—[e]
Yes	2.2 (1.4, 3.6)	2.2 (1.3, 3.5)	—[e]
Weight Gain <15 Pounds			
No	1.0[b]	—[e]	—[e]
Yes	4.6 (2.6, 7.8)	—[e]	—[e]

[a] Analysis confined to Blacks.

[b] Referent.

[c] Index takes into account education (4 categories) and income (5 categories) and places equal percentages of the total Black and White population in three tertiles.

[d] Pre-pregnancy Quetelet Index (wt in pounds/height in inches^2 × 100) ≤2.7.

[e] Variable not included in this model.

TABLE 3—Estimated Relative Risk of Various Sub-Categories of Low Birth Weight by Cocaine Use[a]

		≤1500 g		1501–2500 g <37 weeks		1501–2500 g ≥37 weeks	
Variables	Controls N	Cases N	RR[b] (95% CI)	Cases N	RR[b] (95% CI)	Cases N	RR[b] (95% CI)
Self-Reported Cocaine Use							
Did not use at all	375	78	1.0[c]	116	1.0[c]	91	1.0[c]
Used cocaine and/or crack cocaine, but not throughout pregnancy	5	3	2.9 (0.7, 12.4)	5	3.2 (0.9, 11.3)	4	3.3 (0.9, 12.5)
Used cocaine and/or crack cocaine throughout pregnancy	9	12	6.4 (2.4, 17.2)	20	7.2 (3.0, 17.6)	19	8.7 (3.6, 21.6)

[a] Analysis confined to Blacks: excludes 19 women who could not be classified because gestational age was unknown.

[b] Reference group for this analysis is all Black controls.

[c] Referent.

control" by adjusting for these factors. In multivariate analyses that excluded these variables and adjusted for confounders, the relative risk of having a low birthweight infant in women who used cocaine and/or crack throughout pregnancy was still elevated by a factor of about 4 to 5 (Table 2).

The use of cocaine and/or crack throughout pregnancy increased the relative risk of very low birth weight, pre-term low birth weight, and intrauterine growth retardation in Blacks by about the same amount (Table 3).

Based on an estimate that the prevalence of use of cocaine and/or crack throughout pregnancy is 3.9 percent among Black women in Alameda County and an estimate that the relative risk of low birth weight in users of cocaine and/or crack throughout pregnancy is 4.0 after adjustment for confounding, the attributable risk of low birth weight for cocaine and/or crack use throughout pregnancy in Black women in Alameda County can be estimated to be 0.10.*

Discussion

This study joins a growing body of research[3–11] that links maternal cocaine use with poor pregnancy outcome. The study is unique in being population-based. Results show that cocaine use increases the risk of both pre-term delivery and intrauterine growth retardation, a finding that also is in keeping with the findings of most other similar studies,[5–11] that of Chasnoff, *et al*,[4] being an exception.

This study has some important limitations:

- We did not interview all of the women eligible for the study, although consideration of the characteristics of the non-interviewed women, who mostly could not be located, suggest they would be *more* likely than those we interviewed to use cocaine.
- Information on cocaine use was based on self-report. It is unlikely that women would overreport use of an illegal drug, thus spuriously increasing the risk of low birth weight in association with cocaine use, but we cannot exclude this possibility entirely.
- Mothers of cases of low birth weight might have selectively overreported cocaine use because of concern or guilt about an adverse outcome, or for other reasons. The number of mothers of cases who selectively overreported cocaine use would have to be very large to explain our findings.
- Women may have reported cocaine use but not use of drugs, such as heroin, that are known to reduce birth weight. It is not easy to understand why women would report cocaine use and not use of other illegal drugs, but this explanation for our results cannot be excluded.
- Our study is limited in its ability to draw conclusions about dose-response, because our definition of cocaine use is imprecise. We do not have information about the amount of cocaine use or the timing of its use during pregnancy, because we did not ask these questions.

Considering these limitations and in the context of a large body of data linking cocaine use with intrauterine growth retardation and pre-term delivery,[3–11] we conclude that cocaine use is an important contributor to the public health problem of low birth weight in the Black population of Alameda County, and probably in other urban settings where there is substantial use of the drug. Women who cease cocaine use in the first trimester of pregnancy have been shown to have no increase in the risk of pre-term delivery or growth retardation.[9] Even though specific programs to reduce cocaine use during pregnancy have not been evaluated, the evidence linking cocaine use with poor pregnancy outcome appears sufficient to warrant interventions that provide counseling and a broad-based community effort to identify women early who use cocaine in their pregnancies. The magnitude of the increase in the risk of poor pregnancy outcome in cocaine users suggests that the success of efforts to improve infant and perinatal mortality in settings where use is common may be dependent on addressing the problem of cocaine use successfully.

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Cigarette Smoking as a Risk Factor for Sudden Infant Death Syndrome: A Population-Based Study

BENGT HAGLUND, DRMEDSC, AND SVEN CNATTINGIUS, MD

Abstract: Risk factors for sudden infant death syndrome (SIDS) were examined in a prospective study based on Swedish births between 1983 and 1985. All infants surviving the first week of life were included (279,938). The overall rate of SIDS was 0.7 per 1,000 first week survivors. Elevated relative risks were associated with low maternal age, multiparity, maternal smoking, and male infants. Smoking doubled the risk and a clear dose-response relation by amount smoked was observed. Maternal smoking also seemed to influence the time of death, as infants of smokers died at an earlier age. In countries like Sweden, smoking may be the single most important preventable risk factor for sudden infant death syndrome. (*Am J Public Health* 1990; 80:29–32.)

Introduction

Although syndromes similar to sudden infant death syndrome (SIDS) have been known since biblical time, progress in solving the enigma of SIDS has been slow. The incidence of SIDS varies widely, being low in Scandinavia[1] and higher in other industrialized countries, e.g., in North America. In Sweden[2] and Finland,[3] the incidence of SIDS has doubled since the 1970s, despite reductions in the infant mortality rate of almost 40 percent.[2]

SIDS and other causes of infant death share some risk factors, but the age at death distinguishes SIDS from the other causes of death.[4] SIDS victims usually die within one to six months of age; for unknown reasons, there are two peaks of age at death: eight to nine weeks, and 13 to 15 weeks.[2,5]

Low social class and maternal smoking have previously been associated with SIDS.[1,3,4] Since smoking is more frequent in lower social classes,[6] it is uncertain whether the increased risk for SIDS should be ascribed primarily to smoking or low social class.

In 1973, a medical birth registry was set up by the National Board of Health and Welfare in Sweden,[7] and, since 1982, information has been collected prospectively for each pregnancy from the first antenatal visit. The registry provided an excellent opportunity to investigate the influence of maternal smoking and other risk factors on SIDS in an unselected, homogeneous low-risk population.

Methods

The Swedish National Board of Health and Welfare receives demographic data, previous reproductive history and pregnancy outcome on all hospital births. This is transmitted as copies of the standardized antenatal and obstetrical records, which are generally used in Sweden. In the first trimester, midwives at the antenatal care clinics record previous reproductive history, smoking habits, and family situation. Practically all Swedish women visit these clinics regularly during pregnancy.[8] All births reported to the birth registry are validated every year against a population registry at Statistics Sweden, which receives information from the parishes about all births and deaths. This cross-checking is done by using the mother's personal identification number. The Medical Birth Registry covers more than 99 percent of all births in Sweden.[9]

During the study period from 1983 to 1985, 280,256 infants survived the first week of life. The study was restricted to births of mothers between 15 and 44 years of age. The 318 births for which maternal age was either unknown or outside this interval were excluded, leaving a total of 279,938 infants for study. Ninety-five percent of these infants were born to Scandinavian women.

In Sweden, the death certificate includes information about when the death occurred, possible causes of death, and whether or not an autopsy was performed. In the present investigation, 698 infants were reported to have died between one and 51 weeks of age. SIDS was reported as the only cause of death in 190 infants (International Classification of Diseases, ICD 8, 1968, 795,99). Autopsies were performed on 98 percent of the cases (94 percent forensic and 4 percent clinical). Most deaths from other causes also had autopsies (58 percent clinical and 12 percent forensic).

Parity was defined as the number of live births plus stillbirths. Smoking habits were categorized as non-smoker (i.e., non-daily smoker), moderate smoker (smoking between one and nine cigarettes per day), or heavy smoker (10 cigarettes or more per day). Since it is common in Sweden for young couples to live together without being married, information about family situation was recorded as whether or not living with infant's father.

Multiple logistic regression analyses were performed in order to calculate associations of the independent variables on SIDS as estimated by the relative risk.[10] The independent variables used in the analyses were maternal age, parity, family situation, smoking habits, type of birth (single or multiple), and sex of infant. The continuous variables were categorized in order to avoid assumptions about linearity. The statistical analyses were made using the SAS program package.[11,12]

Results

Mortality from SIDS, 190 cases or 0.7 per 1,000 first week survivors during the three years, accounted for about 27 percent of the total infant mortality after the first week of life. Table 1 shows that maternal age was inversely associated with risk for SIDS. Other characteristics of the mothers associated with an excess number of SIDS were high parity, not cohabiting with infant's father and the number of cigarettes smoked per day. Higher risks for boys and for multiple births were also seen.

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TABLE 1—Number of Infants Surviving the first Week of Life and SIDS Death Rate by Characteristics of the Mother and Infant, Sweden, 1983–85

Characteristics	Infants No.	SIDS No.	SIDS Death-rate per 1000
Maternal Age at Delivery (years)			
15–19	9,574	14	1.46
20–24	66,582	53	0.80
25–29	103,527	68	0.66
30–34	68,838	41	0.60
35–39	27,155	12	0.44
40–44	4,262	2	0.47
Parity			
1	112,994	52	0.46
2	101,323	75	0.74
3	47,333	38	0.80
≥4	18,288	25	1.37
Family Situation			
Living with father	253,588	164	0.65
Not living with father	14,436	17	1.18
Not available	11,914	9	0.76
Smoking Habits			
Non-smoker	181,218	88	0.49
Daily smoker			
1–9 cigarettes	48,465	45	0.93
≥10 cigarettes	31,225	46	1.47
Not available	19,030	11	0.58
Sex of Infant			
Boy	143,981	110	0.76
Girl	135,391	79	0.58
Not available	566	1	...
Type of Birth			
Single	274,526	184	0.67
Multiple	5,407	6	1.11
Not available	5	0	...
Total	279,938	190	0.68

Smoking habits varied with age and family situation: half of the teenage mothers smoked compared with 25 to 29 percent of mothers aged 25 or more. More than half of women who were not living with infant's father were daily smokers, compared to 29 percent of those living with infant's father.[13] We therefore adjusted the estimates for the effects of other variables using a multiple logistic regression model (Table 2).

As seen in Table 2, high parity strongly increased the risk for SIDS, but high maternal age reduced the risk of SIDS. It is plausible that the effect of parity varies with age, but we detected no interaction between age and parity ($p = 0.6$). Infants not living with their father had 50 percent greater SIDS mortality than those living with their father (95% CI = 0.8, 2.5).

Maternal smoking was strongly related to SIDS even when controlling for other risk factors. Smoking up to nine cigarettes per day doubled the risk of SIDS and smoking 10 cigarettes or more per day nearly tripled the risk of SIDS, compared to non-smokers. We examined whether smoking increased the risk of SIDS mortality more for certain women than others, but no clear interactions were detected (smallest p-value = 0.11).

Age at Death of SIDS Infants

The age at death from SIDS ranged widely with a peak incidence at nine to 12 weeks. The age distribution reported from other studies—two peaks during the first six months of life[2,5]—was not clearly observable in our study. Only when the age of death was grouped by three-week intervals was there a weak tendency toward an early and a late peak.

The SIDS victims were stratified into two groups of equal size. However, before this was done, the 10 percent with the longest survival time were excluded in order to obtain groups including only cases with typical ages at death. Early SIDS were defined as deaths occurring between seven and 67 days of age and late SIDS were those between 68 and 145 days of age. Although the magnitude of the risks differed between the two models, they were in general rather similar (Table 2). However, moderate maternal smoking almost tripled the risk for early SIDS, while no increase in the risk for late SIDS appeared. A similar but less marked pattern was true for infants not living with their father.

For further analyses of the difference between early and late SIDS, a logistic regression model based only on the SIDS cases was performed, to estimate the probability of dying before 68 days of age among the SIDS victims. The analysis showed that moderate maternal smoking was strongly associated with increased risk for early SIDS (RR 1.7, 95% CI = 1.2, 2.1). SIDS victims whose mothers were heavy smokers (RR 1.4, 95% CI = 0.9, 1.4), or not living with infants father (RR 1.8, 95% CI = 0.9, 2.3), also had an increased risk for short survival, while infants of mothers 35 years of older seemed to have higher probability for longer survival (RR 0.4, 95% CI = 0.1, 1.1).

The influence of maternal smoking on the time of dying from SIDS, is illustrated in Figure 1. Infants of non-smoking mothers had a symmetric age distribution with a mode value of nine to 12 weeks. Infants of smoking mothers however, had a right-skewed age distribution with a shorter modal time of survival, five to eight weeks.

Discussion

In the present study, the diagnosis of SIDS was based on cases where the death certificate stated SIDS as the only cause of death. As 98 percent of these cases underwent autopsy (and 94 percent forensic autopsy), the diagnostic accuracy in these cases must be considered as sufficient. The diagnostic difficulties of SIDS, when using death certificates, mainly concern underregistration,[2] so the incidence estimated (0.7 per 1000) may be an underestimate.

Maternal smoking has earlier been reported as a risk factor for SIDS.[5,14,15] In the present study, a strong association between maternal smoking and SIDS was found. We have no direct measures of confounders associated with smoking that might adversely affect the infant. For instance, we have no information on whether there were other smokers in the household. However, adjustment for the effects of maternal age, parity, and family situation only marginally reduced the estimates. Such socio-demographic variables are also related to behavioral and lifestyle characteristics that could adversely affect the infant. It therefore seems unlikely that unmeasured confounders would appreciably influence the association. The present results were based on a prospective study, including about 280,000 infants. The population studied was a relatively homogeneous, low-risk population, which further suggests a limited influence of possible confounders.

The information on smoking habits was based on the mothers' reports to the midwives at their first visit to antenatal care. As practically all women in Sweden regularly visit these clinics during pregnancy, the influence of possible selection bias seems negligible.[8] Since pregnant women today are exposed to anti-smoking information, this may have influenced the validity of self-reported smoking habits.

TABLE 2—Relative Risks* of Sudden Infant Death Syndrome Based on a Logistic Regression Model (257,361 Infants Surviving the first Week of Life in Sweden 1983–85)

	All SIDS (1–51 weeks) Relative risk (95% CI)	Early SIDS (7–67 days) Relative risk (95% CI)	Late SIDS (68–145 days) Relative risk (95% CI)
Maternal Age (years)			
15–19	2.0 (1.0, 4.0)	1.5 (0.5, 4.4)	2.6 (0.9, 7.6)
20–29 (ref. group)	1.0	1.0	1.0
30–34	0.7 (0.5, 1.0)	0.6 (0.3, 1.0)	0.8 (0.5, 1.5)
35–44	0.5 (0.2, 0.8)	0.3 (0.1, 0.8)	0.8 (0.4, 1.7)
Parity			
1 (ref. group)	1.0	1.0	1.0
2–3	2.2 (1.5, 3.3)	2.8 (1.6, 4.8)	2.2 (1.2, 3.9)
≥4	4.8 (2.7, 8.4)	5.0 (2.1, 11.9)	5.0 (2.2, 11.4)
Family Situation			
Living with father (ref. group)	1.0	1.0	1.0
Not living with father	1.5 (0.8, 2.5)	1.9 (0.9, 4.0)	0.7 (0.2, 2.2)
Smoking Habits			
Non-smoker (ref. group)	1.0	1.0	1.0
Daily smoker			
1–9 cigarettes	1.8 (1.2, 2.6)	2.8 (1.7, 4.8)	1.0 (0.5, 1.8)
≥10 cigarettes	2.7 (1.9, 3.9)	3.6 (2.1, 6.2)	2.0 (1.1, 3.5)
Sex of Infant			
Boy	1.4 (1.1, 1.9)	1.2 (0.8, 1.9)	1.7 (1.1, 2.7)
Girl (ref. group)	1.0	1.0	1.0
Type of Birth			
Single (ref. group)	1.0	1.0	1.0
Multiple	1.5 (0.7, 3.4)	2.1 (0.8, 5.8)	1.1 (0.3, 4.5)

*Estimates are adjusted for all other variables.

Moreover, women may also have stopped smoking during pregnancy. It has been reported elsewhere that 15 to 20 percent of pregnant smokers quit "on their own" during pregnancy.[16,17] Thus, if a causal relationship between maternal smoking and SIDS exists, the true relative risk may be even greater than estimated herein.

Other studies have reported low birthweight to be associated with SIDS[3,18] and smoking to be associated with low birthweight.[17] Because the effect of maternal smoking on SIDS could be mediated by low birthweight, we also conducted an alternative analysis controlling for low birthweight (<2500 grams). We found almost the same risk estimates for smoking (RR = 1.7 and RR = 2.5 for moderate and heavy smoking, respectively) indicating that smoking affects birthweight-specific as well as overall SIDS mortality.

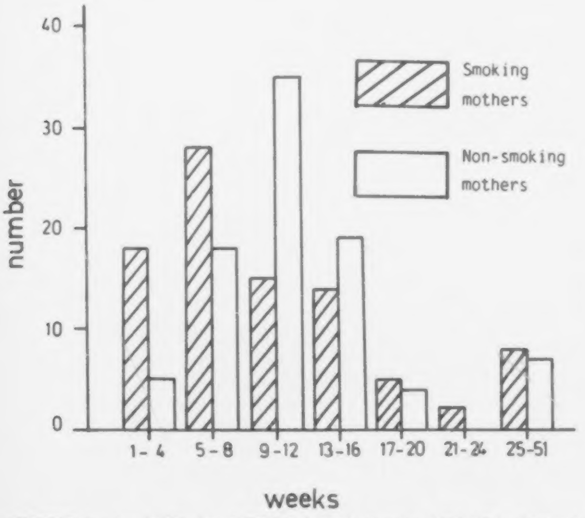

FIGURE 1—Age in Weeks of SIDS Infants in Sweden 1983–85 by Maternal Smoking Habits

The socioeconomic factor included in this study (cohabiting vs not cohabiting parents) has previously been associated with SIDS in Scandinavia,[2,19] but information about maternal smoking was not included in those studies. In the present study, the crude relative risk of SIDS was 1.9 in non-cohabiting families, but daily smoking was more common among non-cohabiting women. Adjusted for the effects of differences in smoking habits, the relative risk was reduced to 1.5 (95% CI = 0.8, 2.5). These data suggest that smoking is a confounding factor, through which at least some of the effects of social class on SIDS rate are mediated, at least in the homogeneous populations of Sweden, where poverty and malnutrition are practically non-existent.

The previously reported bimodal age distribution of SIDS[2,5] led us to stratify the SIDS cases into early and late SIDS. The most striking difference in risk factor pattern between early and late SIDS was the influence of maternal smoking, a finding in accordance with Golding, *et al.*[5] Early SIDS was more strongly related to maternal smoking. We suggest that the reported bimodal age distribution may partly reflect differences in maternal smoking.

Today, smoking is probably the most important preventable risk factor for late fetal death.[13,20] The present results suggest that this may also hold true for SIDS. Assuming that maternal smoking could be eliminated and infants delivered by mothers who gave up smoking faced the same mortality risk as their non-smoking peers (i.e., same maternal age, parity, sex, etc.), we estimated that SIDS death rate would be reduced by 27 percent. In terms of number of SIDS deaths, this means a reduction from 190 deaths to approximately 140 deaths during a three-year period in Sweden.

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Past Pregnancy Outcomes among Women Living in the Vicinity of a Lead Smelter in Kosovo, Yugoslavia

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Abstract: This analysis compares the rates of spontaneous abortion among women living in the vicinity of a lead smelter with those of women living in a town where blood lead levels were low. Data derive from the obstetric histories of both groups of women obtained while seeking prenatal care for a later pregnancy. A total of 639 women (304 exposed, 335 unexposed) had at least one previous pregnancy and lived at the same address since their first pregnancy. The geometric mean blood lead concentrations in the sample at the time of the interviews were 0.77 μmol/L in the exposed town and 0.25 μmol/L in the unexposed town.

The rates of spontaneous abortions in first pregnancies were similar, with 16.4 percent of women in the exposed town and 14.0 percent in the unexposed town reporting loss. The adjusted odds ratio relating town of residence to spontaneous abortion was 1.1 (95% CI = 0.9, 1.4). This analysis represents the first systematic attempt to seek an association between environmental lead exposure and spontaneous abortion. As such, the failure to find a positive association strongly suggests that at the levels of exposure represented in our sample, such an association does not exist. (*Am J Public Health* 1990; 80:33–35.)

Introduction

Historical anecdotes have suggested that occupational exposure to lead may be detrimental to reproductive outcome. Pregnant women working in the pottery and white lead industries in England in the nineteenth century were reported to be more likely to miscarry than women not so employed; other reports from Italy, Germany, France, and Japan also describe high rates of loss among pregnant women exposed to lead as compared to women without exposure.[1] While most samples lack both precise measures of exposure and information on potentially confounding factors, taken together they suggest that high levels of lead might be harmful to the developing conceptus. This suggestion is consistent with a series of reports from Sweden, where Nordstrom and co-workers found a significantly increased frequency of spontaneous abortion among female employees working at a lead smelter.[2] Similar observations were made on populations living in close proximity to that smelter.[3]

In contrast, a recent prospective cohort study of pregnant women living near a lead smelter in Australia reported no association between maternal blood lead (BPb) and spontaneous abortion.[4] That study, however, may have suffered from incomplete ascertainment of the event.[5] The present report examines the rates of spontaneous abortion in two comparable populations, one living in the vicinity of a lead smelter and the other in a town 25 miles away without significant environmental lead exposure.

Methods

The subjects of this analysis were recruited for a prospective cohort study of environmental lead exposure, reproductive outcome, and infant development. Further description of the population and methods of data collection have been presented in a recent publication.[6] The past reproductive histories of these women provide the data for this report. In brief, the study took place in Titova Mitrovica (T. Mitrovica), Yugoslavia, where there is significant environmental exposure to lead from a lead smelter, refinery, and battery plant, and in Pristina, Yugoslavia, a town without significant lead exposure. The prospective cohort study recruited women in mid-pregnancy. At that time an interview inquiring about medical and reproductive history and sociodemographic factors was administered.

A total of 1,502 women was recruited at mid-pregnancy. This analysis draws on data from the first 1,032 recruited during the period of May 1985–May 1986. Of the 1,032 women evaluated, 777 (332 in T. Mitrovica and 445 in Pristina) had had at least one previous pregnancy. In T. Mitrovica, the current BPb varies with distance from the smelter, such that women living close to the smelter have, on average, higher blood lead (BPb) levels than those living further away.[6] Since it is reasonable to assume that this variation in BPbs with distance from the smelter was also true in the past, it is only among women who have not moved since their first pregnancy that we can be assured of previous lead exposure during that pregnancy. Accordingly, this analysis was confined to 639 women (304 in T. Mitrovica and 335 in Pristina) who had lived at their current address at least since their first pregnancy.

In these analyses, spontaneous abortion was defined as a fetal loss prior to the seventh month of gestation; stillbirth was defined as a fetal loss occurring in months seven through nine of gestation. Only the first pregnancy of each woman was used in the analysis. This strategy avoids the problem of nonindependence of reproductive outcomes within individuals.[7,8] Maximum likelihood logistic regression analysis was used to assess the influence of lead on spontaneous abortion, after adjustment for the effects of potentially confounding variables.[9]

Results

Table 1 shows the demographic characteristics of the 639 women with at least one previous pregnancy who lived in the same home since their first pregnancy. Mean parental ages at

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TABLE 1—Demographic Characteristics at Study Recruitment

Characteristics	T. Mitrovica	Pristina
Number of women	304	335
Mean maternal age at first pregnancy	21.7 (3.9)*	22.1 (3.4)
Mean paternal age at first pregnancy	26.0 (4.8)	25.6 (4.0)
Mean maternal education (years)	8.7 (4.0)	8.7 (3.9)
Mean paternal education (years)	11.3 (2.7)	11.2 (3.0)
Mean number of previous pregnancies	2.8 (2.2)	2.6 (2.0)
% Current cigarette smokers	23.0	26.3
% Current alcohol users	5.9	10.5
% Same address since birth	31.0	30.2
Mean years at current address	13.6 (10.1)	7.0 (6.4)
% Husbands employed in Pb industry	28.8	—
Ethnicity		
% Albanian	59.3	63.3
% Serbian	21.5	25.6
% Other	19.2	11.1

*Standard deviation.

first pregnancy, mean parental years of school at the later index pregnancy, and gravidity were similar in the two towns. Women in Pristina compared to women in T. Mitrovica were slightly more likely to be current smokers and twice as likely to be current alcohol users. While approximately 30 percent of women had lived in their respective towns since birth, women in T. Mitrovica had lived an average of six years longer at their current address than women in Pristina. Within this sample, the proportions of women of Albanian and Serbian origin were higher in Pristina.

Hematologic findings at the time of study recruitment revealed the expected differences in BPb and erythrocyte protoporphyrin (EP) (Table 2). Mean BPb in T. Mitrovica was 0.77 μmol/L as compared to 0.25 μmol/L in Pristina. The mean hemoglobin concentrations were within normal limits in both towns.

The rates of spontaneous abortions were similar in T. Mitrovica (16.4 percent) and Pristina (14.0 percent) (Table 3). Stillbirth rates were also similar across towns. When all fetal losses were combined, the rate in T. Mitrovica was 19.7 percent; the rate in Pristina was 17.9 percent. These rates remained essentially unchanged when adjusted for the number of induced abortions in each town.[10]

Adjustment for the effects of maternal age at first pregnancy, current smoking, ethnic group, and maternal education by maximum likelihood logistic regression produced little change in the odds ratio relating town of residence and spontaneous abortion (odds ratio 1.1) (Table 3). Refinement of the exposure variable using distance of the home from the smelter showed no consistent trend in the odds ratio using Pristina as the reference group. For women living within 2 km of the smelter, the adjusted odds ratio was 1.2 (95% confidence interval = 0.7, 1.9); for those within 2–4 km it was 2.1 (0.8, 5.3); for those within 4–8 km it was 0.5 (0.2, 1.2); and for those beyond 8 km it was 0.8 (0.6, 1.6). Analysis of the data restricted to women who reported never drinking alcohol showed similar results.

TABLE 2—Current Blood Lead and Hematology Findings

	T. Mitrovica	Pristina
Number of women	304	335
Blood Lead (μmol/L)[a]	0.77 (0.32, 1.83)	0.25 (0.12, 0.51)
Hemoglobin (g/L)[b]	125 (103, 147)	123 (94, 151)
EP (μmol/L)[a]	0.71 (0.21, 2.41)	0.50 (0.22, 1.17)
Serum Ferritin (μg/L)[a]	12.5 (1.9, 82.9)	10.8 (1.4, 85.0)

[a] Geometric means (95% confidence interval)
[b] Arithmetic means (95% confidence interval)

TABLE 3—First Pregnancy Outcome Rates

	T. Mitrovica	Pristina	Odds Ratio[c]
	% (number)	% (number)	(95% CI)
Spontaneous Abortions[a]	16.4 (50)	14.0 (47)	1.1 (0.9, 1.4)
Stillbirths[b]	3.9 (10)	4.5 (13)	1.0 (0.6, 1.5)
Livebirths	80.3 (244)	82.1 (275)	—

[a] Denominator excludes reported induced abortions, ectopic pregnancies, and pregnancies where the outcome was unknown.
[b] Denominator excludes spontaneous abortions.
[c] Adjusted for maternal age, current smoking, ethnic group, and maternal education.

Discussion

These data indicate that lead exposure, at the levels encountered in T. Mitrovica, is not a risk factor for spontaneous abortion. The adjusted odds ratio was 1.1 with an upper bound of 1.4 suggesting larger effects are unlikely. There was 80 percent power to detect an odds ratio of 1.6 (alpha = 0.05, two tailed).

In order to increase our confidence that this result is real and not due to bias, we examined carefully the basis for measuring exposure, outcome, and confounding variables. Town of current residence is the index of exposure in this analysis and the population of women examined lived at least since their first pregnancy in the same home. While current blood lead is a marker of recent exposure, the striking differences in current blood leads between these two groups, together with the fact that these women had not moved since their first pregnancies, make it likely that exposure at first pregnancy similarly differed across towns. Moreover, previous blood and air lead survey data from these two towns indicate that the levels of exposure in T. Mitrovica were probably higher in the past.[11] Thus, previous levels of lead exposure such as might have occurred in the first pregnancies of these study women are likely to have been higher than those recorded at present. We remain assured that the index of exposure is an appropriate one, that the exposure differs across towns, and that the levels of exposure in T. Mitrovica among currently pregnant women are significantly higher than those in Pristina. We note, however, that the level of lead exposure of our subjects in T. Mitrovica does not nearly approach that typically found among women exposed to lead through occupation.

In reviewing this study, one could speculate about the possible effects of random or differential misclassification of spontaneous abortion. Insofar as spontaneous abortion rates are similar to those reported in other samples, the stillbirth rates comparable to those found in less developed countries, there is no basis for positing high levels of misclassification. Differential misclassification might be produced if recall differed between towns. In order to produce the null result observed here, we must posit bias in opposite directions for the two towns: underreporting in T. Mitrovica and overreporting in Pristina. Again, since the rates of loss in both towns fall well within the bounds usually reported, such biasing effects seem unlikely.

Previous work which examined the relation of lead exposure to spontaneous abortion revealed inconsistent findings. The initial historic reports which suggested an association are largely based on anecdotal reports of women working in lead-related industries where more severe and multiple occupational exposures may have been operating. These observations gave rise to the hypothesis of a causal relation, but did not provide systematic evidence of such an association. The Swedish smelter studies,[2,3] although suggestive of an effect of environmental exposure on spontaneous abortion, were largely ecologic and thus do not provide strong support for an association at the individual level.

The only prospective cohort study to address the issue of lead and reproductive outcome assessed 831 pregnant women living near a lead smelter in Australia.[4] In that study, mean blood lead levels (std error) among the exposed were relatively low, ranging from 0.51 ± 0.01 to 0.54 ± 0.01 μmol/L. Blood lead and other risk factors were examined in relation to pregnancy outcome and no association was found between maternal BPb and spontaneous abortion. However, the number of abortions ascertained was small ($n = 23$) raising the question of incomplete ascertainment of the event.[5] The failure to find a positive association in the current study strongly suggests that, at the levels of exposure represented in our sample, such an association does not exist.

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Call for Proposals: Improving Service Systems for the Physically Disabled

The largest single commitment of private funds ever provided to benefit people with disabilities has been announced by the Robert Wood Johnson Foundation of Princeton, NJ. The Foundation will award up to $8.4 million to community-based organizations run by and for the disabled. As many as 12 one-year planning grants of up to $100,000 will be made under the Foundation's new program, "Improving Service Systems for People with Disabilities."

The grant funds are intended to help community organizations such as Independent Living Centers create comprehensive service systems so that the disabled can live independently. Community organizations that receive the planning grants also will be eligible to compete for three-year grants of up to $600,000 to carry out their plans to coordinate available health and supportive services, develop new services, and identify public and private financing sources for these services.

The application deadline is January 16, 1990. Organizations intending to apply should contact Lex Frieden, Independent Living Research Utilization, 3400 Bissonnet, Suite 101, Houston, Texas, 77005 or call 713/528-0504. The grants awards will be announced in the fall of 1990.

The Effect of a Patient Education Program on Emergency Room Use for Inner-City Children with Asthma

MARK C. SHIELDS, MD, MBA, KENNETH W. GRIFFIN, BA, AND WYLIE L. MCNABB, EDD

Abstract: An educational program for children with asthma designed to reduce emergency room (ER) use enrolled all eligible children (n = 253 primarily low-income Black) within a health maintenance organization (HMO) who had used the hospital or ER for asthma during the pre-enrollment period and randomized them into two groups. Twenty-four of the experimental group patients had 55 ER visits and 18 of the control patients had 39 ER visits during the first 12 months post-intervention. This program did not achieve its goal. (*Am J Public Health* 1990; 80:36–38.)

Introduction

Asthma is a common disease that leads to considerable morbidity and expense.[1] Lethal asthma is a serious problem and continues to increase in frequency, especially for minority patients.[2] The years of potential life lost due to asthma almost doubled for minority males between 1979 and 1983.[3] One of the strategies to reduce morbidity, mortality, and expense of asthma has been patient education. The program evaluated in this paper was designed to improve patients' knowledge and behaviors related to asthma in order to improve their control of the disease, and thereby reduce emergency room (ER) use.

Methods

Subjects

Subjects were drawn from two facilities of an urban health maintenance organization (HMO) with a combined enrollment of over 50,000. Approximately 80 percent of these patients were Black and all had at least one working adult in the family. The median household income for the HMO service area was 6 percent below the overall city median. Patients were eligible for inclusion in the study if they were 18 years of age or younger and had at least one emergency room visit or hospitalization for asthma during the prior four-year period. A total of 253 subjects met eligibility requirements. All eligible patients were enrolled in the study.

Randomization

For randomization purposes, patients were listed in order of their past emergency room use and descending this list patients were allocated randomly using a table of random numbers to either experimental (n = 127) or control (n = 126) groups. The random assignment was balanced so that out of every 10 assignments, five were to the experimental group. The experimental and control groups were virtually identical on several measures as shown in Table 1.

TABLE 1—Profile of total Group [N = 253]

	Experimental Group	Control Group
Number of Patients	127	126
Mean Age (years)	8.5	9.0
Sex (% male)	61%	65%
ER Use before entry (# visits in 1 year)	1.2	1.2
Months in HMO before entry	20.5	21.2
Percent registered at largest HMO center	48.0	49.2
Percent with one or more hospital days for asthma in prior year†	18.2	23.6

†Percent shown for patients for whom data are available.

Data Collection

The HMO collected information on hospital days, emergency room visits, and office visits including the related diagnoses during its normal course of business. Once a patient dropped out of the HMO, no further information was available. The enrollment and dropout of patients were monitored by the HMO. Utilization outside the HMO while a member was recorded because virtually all other emergency rooms and hospitals would have billed the HMO for their services.

Educational Program

The program consisted of two components: classes conducted in group settings for patients and their parents, and telephone instruction by protocol to patients or parents by nurse clinicians to provide additional instruction and reinforcement. The classes and calls covered 23 instructional objectives grouped into four content areas: 1) prevention of asthma attacks, 2) medication management, 3) intervention during asthma attacks, and 4) utilization of health care resources. Instruction dealt with knowledge, skills, and behavior. Four classes were offered to each patient and his or her parents on weekday evenings and Saturday mornings with make-up sessions scheduled periodically to accommodate patients who missed classes. Each class lasted 1.5 hours. Following classes, telephone instruction by protocol to patients and their parents was done at least each season. The calls typically lasted 30 minutes.

The program was offered to all 127 experimental group patients. Of the patients who remained in the HMO for 12 or more months, 38 percent attended one or more classes. Telephone instruction by protocol improved patient contact to 81 percent. The mean number of educational contacts (classes plus calls) was 4.4. The intervention increased the frequency of self-reported mastery on most of the instructional objectives. The post-intervention mastery of specific educational objectives varied from 80 percent of patients for avoidance of over-the-counter drugs, 75 percent of patients for quitting smoking, 35 percent of patients for recognition of

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the early signs of an asthma attack, to 30 percent of patients for using medicines correctly. Detailed descriptions of the program content, instructional techniques, and educational outcomes for these patients have been reported previously.[4,5]

Exposure of control group patients to the intervention was unlikely because the program was performed apart from the primary care providers and the program staff avoided routine contact with those providers. The control group continued to receive their usual medical care.

Results

Of the 253 patients who began the program, 205 remained within the HMO for the first 12 months and 194 remained within the HMO for the entire study period, up to 29 months. Dropouts were evenly distributed between experimental and control groups. Participants and dropouts were compared according to age, sex, and ER use before entry. The differences were small and not statistically significant.

There was no significant difference between experimental and control groups in post-intervention health care utilization for respiratory related illnesses for a one-year period as shown in Table 2. This analysis excludes dropouts and eliminates any seasonal bias, since each patient was enrolled for 12 continuous months. Hospital days and office visits also are compared. Further analysis (not shown in this table) revealed no differences in utilization for non-respiratory illnesses. The results were similar when this analysis was repeated for the total time period of the study with all dropouts included. The probability of a Type II error for a clinically significant difference was calculated and found to be negligible.

Discussion and Conclusions

The program did not achieve its goal of decreasing emergency room visits for children with asthma. This occurred despite careful program design, implementation by clinicians with expertise in pulmonary disease, and an average of $96 expended per patient per year (1984 dollars) for the asthma education.

Other authors have shown other programs to be effective with children with asthma through such outcome measures as reduced emergency room visits, hospital use, sick days, and days lost from school.[6–10] Why have these other interventions been successful while this program was not? The inability of this intervention to demonstrate a reduction in ER utilization does not appear to be due to any of the following methodological problems: randomization problems, dropouts, exposure of control patients to intervention, omission of traditional program content, or accuracy of outcome measures.

TABLE 2—Respiratory-Related Utilization for Patients Enrolled Entire 1st 12 Months [N = 205]

	Experimental Group	Control Group
Number of Patients	101	104
Patient months enrolled	1212	1248
Number of ER visits	55	39
Distribution of ER visits		
0	77	86
1–2	17	13
3 or more	7	5
Hospital Days		
mean	0.27	0.22
standard deviation	1.34	0.74
Office Visits		
mean	1.63	1.86
standard deviation	2.28	4.24

The program probably failed for two sets of reasons: failures of implementation, and failures of program effectiveness. The socioeconomic, educational, and racial profile of the patients probably contributed to problems with both program implementation and program effectiveness. Programs need to be tailored to their target populations. Furthermore, the program was difficult to implement because it was offered to all eligible children with asthma, not just volunteers. This undoubtedly accounted for some of the difficulties in obtaining participation in classes by patients and parents. Implementation was further hindered by the enrollment of many patients with moderate asthma who were probably less motivated and less likely to benefit than patients with more severe disease.[11]

Problems with program effectiveness took several forms. The behavioral component of this program was less intensive than that of other successful programs.[6] It has been demonstrated for asthma[12] and other diseases[13–15] that increases in patient knowledge alone, without instruction on how to change behavior, will not improve patient compliance or health outcomes. Although the importance of behavioral change was discussed with patients in this program, specific behavior modification techniques such as reinforcement and contingency contracting were not utilized. Secondly, less mastery of the educational material was attained by the patients than anticipated, even though considerable resources were used in the program. Less mastery occurred partially because class attendance was less than expected. Furthermore, although telephone instruction alone improved the rate of contact and increased the level of mastery of certain educational objectives, it did not lead to the level of mastery desired.[5] Thirdly, those instructional sessions with parents alone may not have been as effective as those sessions including the children.

Although this study did not show any impact on health care utilization, it is possible that the intervention had a positive impact on functional measures such as days lost from school or sick days. These measures should be included in future research.

This study identifies problems that must be addressed during the widespread implementation and funding of patient education for asthmatic patients. To the inexperienced, patient education seems deceptively simple. However, unless the complexity of patient education is appreciated, a program may fail and may have unintended outcomes. The findings of this study suggest that programs should: 1) include a strong behavioral component, 2) be targeted at key groups who will actively participate and benefit, 3) be monitored to assure that intermediate goals are reached, and 4) be validated for specific socioeconomic groups before implementation.

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New NCHS Report Gives Trends in Low Birthweight

The National Center for Health Statistics recently released its latest findings in vital and health statistics. The report—*Trends in Low Birthweight: United States, 1975–85*—analyzes national and regional trends in rates of low, moderately low, and very low birthweight, and examines the relative risk by maternal characteristics such as age, race, education, and marital status. Some of the findings:

- Low birthweight (LBW) which had been on the decline for the past two decades, increased 2 percent between 1985–87.
- Approximately 262,000 LBW babies were born in 1987 (69 per 1000 live births).
- Rates of LBW and very low birthweight (VLBW) are substantially higher for Black than for White births. In 1987, Blacks were almost three times as likely as Whites to have a VLBW infant and more than twice as likely to have an LBW baby.
- Rates of low, very low, and moderately low birthweight (MLBW) are lowest for mothers in their late twenties and rise as maternal age increases or decreases.
- All LBW rates are substantially higher for young women having a second or higher order birth than for first-time mothers, reflecting the close spacing of many second and subsequent births.
- More than half of all LBW and MLBW babies and 90 percent of VLBW babies are born before 37 weeks of gestation (preterm). Of LBW babies who were full term, only 3 percent weighed less than 1500 grams, but 27 percent of those born preterm weighed this little. For both preterm and full-term LBW babies, the proportion weighing less than 1500 grams is higher for Black than for White babies.
- For all characteristics examined, rates of LBW, MLBW, and VLBW are substantially higher for Black than for White births. The racial differential is most pronounced for VLBW. In 1985, the VLBW rate was 2.8 times as high for Black as for White babies, and rates of LBW and MLBW were 2.1–2.2 times as high for Blacks as for White births.

The NCHS report, by Selma Taffel, is: *Trends in Low Birthweight: United States, 1975–85*, Vital Health Statistics Series 21, No. 48, DHHS Pub. No. (PHS) 89-1926. For further information about this report, or for a list of titles of reports published in the Vital and Health Statistics series, contact: Scientific and Technical Information Branch, NCHS, Hyattsville, MD 20782, tel: (303) 436-8500.

Post-Hospital Support Program for the Frail Elderly and Their Caregivers: A Quasi-Experimental Evaluation

Julianne S. Oktay, MSW, PhD, and Patricia J. Volland, MSW, MBA

Abstract: This paper evaluates the Post-Hospital Support Program for the frail elderly and their caregivers using a quasi-experimental design. The program goals were to reduce stress in the caregivers, improve functioning and reduce mortality in patients, and reduce health service utilization in patients. Subjects were patients at the Johns Hopkins Hospital in Baltimore, Maryland, age 65 or over, who were returning home to the care of a non-paid caregiver and who had extensive post-hospital care needs which were expected to continue for at least one year. The 93 comparison group patient/caregiver pairs were discharged between May 15, 1983 and May 14, 1984. The 98 treatment group pairs were discharged between May 15, 1984 and May 14, 1985. Interviews were conducted with patients and caregivers at 1, 3, 6, 9, and 12 months after discharge. Results showed a slight reduction in caregiver stress and a substantial reduction in hospital days used by the treatment group. When confounding and history effects are taken into consideration, an average difference of 6.5 days per patient remains. The data also suggest that the treatment program may have postponed some deaths and nursing home placements. The results suggest that support services for the frail elderly and their caregivers can be cost efficient by reducing hospital length-of-stay. Savings were estimated at $4,585 per patient per year in this study. However, further work is needed to design programs which more effectively reduce caregiver stress. (*Am J Public Health* 1990; 80:39–46.)

Introduction

The elderly population is expanding rapidly and, at the same time, growing proportionately older.[1] This means more and more elderly will be in need of some type of long-term care.[2] The cost for long-term care, especially for institutional care, has been growing considerably faster than the rate of inflation.[3] The $33 billion average annual expenditure (1986–90) for nursing home care is expected to expand to over $120 billion by the year 2020.[4] Nor are the projected increases limited to nursing home care. For every frail elderly person now in a nursing home, there are two equally frail elderly living in the community.[5] Providing home care services for this population is increasingly recognized as an important goal while, at the same time, the potentially explosive costs are greatly feared. A key to achieving good care for the community-based frail elderly population, and avoiding costly institutional care, is the family caregiver, who currently provides 80–90 percent of in-home care in the United States. The Informal Caregivers Survey, conducted as a part of the 1982 Long-Term Care Survey, showed that the typical caregiver of the frail elderly is a 57-year-old married female who is in poor health;[6] substantial proportions experience emotional distress and social isolation.[7,8] A number of programs designed to support these family caregivers have sprung up around the country, but most of them offer no more than a brief "support group"; these programs have not been systematically evaluated.[9] We do not yet know whether a program designed to provide support for the family caregiver can be effective in reducing the negative effects of caregiving, and whether this strategy will lead to a reduction in the use of institutional care. This paper describes a study designed to determine the effectiveness of a support program for caregivers of the frail elderly following hospital discharge. The program goals were to reduce stress in the caregivers, improve functioning in patients, reduce mortality in patients, and reduce cost by reducing health service utilization.

Method

Treatment Program

The Post-Hospital Support Program provided both medical and social services to the frail elderly and their families for one year following hospitalization. A coordinated approach was provided by a project team made up of a nurse and a social worker. Services included assessment, case management, skilled nursing, counseling, referrals, respite, education, support group, medical back-up, and on-call help. All cases received a minimum of one nurse and one social work visit per month. However, the amount of service depended on need. The service differed from what is usually available in the Baltimore area not only in the amount of service, but also in its coordinated medical-social approach. Focus was on the caregiver/patient configuration and not primarily on the patient; service was closely coordinated with the hospital inpatient and outpatient services; and service extended for one year post-discharge, regardless of the health status of the patient. Additional services were available, such as respite, some "gap filling" money (used primarily for incontinence supplies), and a support group. A more detailed description of the treatment program is provided in Appendix A.

Subjects

Subjects were inpatients at the Johns Hopkins Hospital who gave informed consent and who met the following eligibility criteria:

- Post-hospital care would be provided by a non-paid caregiver.
- Post-hospital needs were comparable to those of patients in "skilled" nursing facilities. (A measure, used by the State of New York—the DSM-1[10]—to determine eligibility for "skilled" care, was used to determine level of need. Patients with scores of 180 or more were eligible for the study.)
- Post-hospital care needs were chronic (expected to last at least one year) and long-lasting (patient expected to live at least one year).

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- Patient residence was within 20 miles of the Hospital.

Design and Procedures

The design was quasi-experimental.[11] Patients who were discharged from the hospital between May 15, 1983 and May 14, 1984 constituted the comparison group; 93 patients and caregivers who met the eligibility criteria and agreed to participate during the first year of the study formed the comparison group. Patients discharged between May 15, 1984 and May 14, 1985 constituted the treatment group; 98 cases who met the eligibility criteria and agreed to participate in the second year formed the treatment group. The refusal rate was about 10 percent for both groups. Both patients and caregivers were interviewed within one month of discharge, and again at 3, 6, 9, and 12 months post-discharge. At the final interview, there were 50 patient/caregiver pairs in the comparison group and 62 patient/caregiver pairs in the treatment group. The interviews, which were identical for the two groups, were conducted by trained interviewers. Interviewers attended weekly supervisory sessions. Although there was no way to blind the interviewers, they were not aware of the pattern of the design, and did not seem to differentiate between comparison and treatment groups.

Measurement

Caregiver Stress—Caregiver stress was measured in three ways: The 27-item General Health Questionnaire (sensitivity = 88 percent, specificity = 84 percent)[12] was used to measure overall stress level. We also used a symptom checklist for measuring physical health (alpha = 82, inter-rater reliability = .88).[13] A measure of the extent to which caregiving was having a negative impact on the social functioning of the caregiver (Negative Impact of Caregivers Scale) was also included.[14] Finally, caregiver utilization of health services was measured with caregiver report.

Patient Functioning—Patient functioning was measured with the Katz Activities of Daily Living Scale (ADL)[15] (coefficient of reproducibility = .94–.98), three items* from the Instrumental Activities of Daily Living Scale (IADL)[16] (reproductibility = .94) and the Mental Status Questionnaire (MSQ)[17] (alpha = .84, test-retest reliability = .8). The ADL and IADL were based on caregiver report. The MSQ was given directly to the patient.

Patient Utilization of Services—At the 3, 6, 9 and 12 month interviews, utilization of a wide range of health services was determined (by caregiver report) including: physician and emergency room visits, nursing home days, use of home health nurse, home health aide, respite, transportation, social work, and other in-home services. In addition, hospital records were examined to determine hospital days used by the patient.

Data Analysis

Caregiver stress and patient functioning were measured with proportional change scores in each of the four time periods (0–3 months, 3–6 months, 6–9 months, and 9–12 months). That is, the time one score was subtracted from the time two score. This difference was then divided by the maximum change possible given the time one score. Thus if a patient scored "8" on the Mental Status Questionnaire, the maximum positive change would have been "2" because the maximum score on the scale is "10". If the change was negative, it was divided by the maximum negative change possible—in this case "−8". The resulting score ranges from −1 to +1. A score of "0" represents no change. "+.50" represents an improvement of 50% of the possible improvement, given the time-one score. The use of a proportional change score compensates for different amounts of possible change based on any uneven distribution of initial scores. All cases who were available at the two time points were included in the analysis. Thus a patient who died seven months after discharge would be included in the 0–3 and 3–6 month measures but not in the 6–9 or 9–12 month ones. Thus the number of cases decreases over the four time periods. A 0–12 month figure is also included for those who remained in the program for 12 months.

Mortality data are analyzed using the Mantel-Haenszel odds ratio for each of the four time periods and the total 12-month period. A Kaplan-Meier survival curve is also presented. In analyzing service utilization, descriptive measures (e.g., percents, means) are used for each of the time periods. For hospital days and outpatient visits, means are calculated for the entire group and again for the patients who stayed in the study for 12 months. This is to eliminate the effect of any differential mortality or institutional placement rates in the two groups.

Results

Sample Characteristics

The 191 patients in the study were predominantly female (61 percent) Black (76 percent) widowed (53 percent). The mean age was 76 years. Mean annual income was $3,880. One-third were on Medicaid; 63 percent had less than eight years of education. The mean DMS-1 score for the study population was 320. An example of a patient with this score would be Mr. L., "a 71-year-old man with coronary artery disease, diabetes mellitus, hypertension, permanent pacemaker, and glaucoma. He was admitted to the study after being hospitalized for left femoral peroneal bypass. He had extensive wounds extending from the groin to the achilles area requiring daily soaks, dressings and monitoring." In spite of this level of dependency, only 10 percent had poor prognosis; 52 percent had "guarded" or "fair" prognosis, and 38 percent had a good prognosis. The caregivers, 80 percent of whom were female, had a mean age of 53; 30 percent were spouses of the patient, and 35 percent were daughters; 73 percent had been caring for the patient before the hospital episode in which the patient was identified for the study.

Comparison of Treatment and Control Groups

The two groups were similar on most patient and caregiver demographic characteristics and on medical and functional status (Tables 1 & 2), with two exceptions. In the treatment group, more caregivers were new to the task, and fewer patients had a "good" prognosis.

Comparison of Services Received

In the Baltimore area, a variety of services were available to the comparison group during the course of the study. For example, home health service (limited to the period of skilled nursing need) including visiting nurse, home health aide, and social work services were widely available (Table 3). There was also a case management service (continuation of the program evaluated in the National Channeling Demonstration).[18] Transportation was available for Medicaid patients only. A comparison of the rates of service utilization

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*The items were: handling the telephone, handling own medication, and handling own finances.

TABLE 1—Demographic Characteristics of Patients and Caregivers in the Comparison and Treatment Groups

Demographic Characteristics of Patients	Comparison Group (N = 93)	Treatment Group (N = 98)
% female	60.2	61.2
% Black	78.5	74.5
% widowed	54.8	52.0
% <8th grade education	64.8	61.3
% Protestant religion	76.3	74.5
Mean age (years)	76.5	76.3
Mean income*	$5,436.74	$5,747.88
% lives alone	9.8	9.3
% on Medicaid	37.4	28.4
Demographic Characteristics of Caregivers		
% female	82.8	77.6
% of caregivers who are the spouse of patient	28.6	31.9
% of cases where caregiver provided care prior to the index hospitalization	60.4	83.7†

*Adjusted for inflation rate between 1983 and 1984 of 4.3 percent.

†Difference (60.4–83.7) = 23.3, 95% confidence intervals of the difference −35, −12.

TABLE 2—Medical and Funcational Characteristics of Patients in the Comparison and Treatment Groups

	Comparison Group (N = 93)	Treatment Group (N = 98)
Mean DSM-1, score	305.7	333.4
% of "Good" prognosis	46.2	29.6*
% Taking four or more medications	56.1	55.1
Mean ADL score on initial interview	5.0	4.3
Mean MSQ score on initial interview	5.5	6.1
Mean IADL score on initial interview	1.4	1.2
Mean # hospital days in year preceding	7.2	9.5

*Difference (46.2–29.6) = 16.6, 95% confidence intervals 3.6, 29.6.

TABLE 3—In-Home Service Utilization by Comparison and Treatment Groups at 3, 6, 9 and 12 Months Post Discharge

	Comparison Group		Treatment Group		% Difference	95% Confidence Interval
	N	% (a)	N	% (b)	(a − b)	
Visiting Nurse						
0–3 months	66	51	81	96	−45	−31, −59
3–6 months	51	28	73	95	−67	−50, −84
6–9 months	50	28	69	97	−69	−52, −86
9–12 months	50	16	62	95	−79	−61, −97
Home Health Aide						
0–3 months	66	41	81	48	−7	+10, −24
3–6 months	51	25	73	45	−20	−3, −37
6–9 months	50	25	69	46	−21	−3, −39
9–12 months	50	24	62	42	−18	0, −.36
Social Worker						
0–3 months	66	23	81	93	−70	−54, −86
3–6 months	51	12	73	96	−84	−67, −100
6–9 months	50	11	69	91	−80	−62, −98
9–12 months	50	14	62	81	−67	−48, −86
Transportation						
0–3 months	66	30	81	40	−10	+6, −26
3–6 months	51	23	73	43	−20	−3, −37
6–9 months	50	26	69	49	−23	−5, −41
9–12 months	50	28	62	48	−20	−1, −39
Respite Care						
0–3 months	66	14	81	12	+2	+13, −9
3–6 months	51	12	73	15	−3	+96, −15
6–9 months	50	9	69	6	+3	+13, −7
9–12 months	50	14	62	6	+8	+19, −36

in the comparison and treatment groups by caregiver report shows that, in all time periods, more patient/caregiver pairs in the treatment group received visiting nurse service and social work service, home health aids and transportation from the third to the twelfth month post-discharge, but not more respite care.

Impact on Caregiver Stress

A comparison of the proportional difference scores on caregiver stress for the treatment and comparison groups is shown in Table 4. Most of the proportional difference scores are negative, indicating a reduction in stress over time. The exception is on the Negative Impact on Caregiving Scale, which shows a sharp reduction in the 0–3 month period followed by a sharp increase in the 9–12 month period. Overall, the differences between the two groups are very small. The confidence interval includes zero for all but two comparisons: Number of Caregiver Illnesses (6–9 months) and Negative Impact of Caregiving (0–3 months). There is a consistent pattern of greater reduction in the General Health Questionnaire scores (indicating less mental disorder) for treatment group caregivers, but none is large enough to inspire confidence.

No clear relations were found between prognosis and the caregiver stress variables. However, previous caregiving experience was found to be related to the proportional difference in caregiver illness at 6–9 months ($r^2 = .06$). (Those with prior caregiving had greater reductions in numbers of illnesses.) A pooled proportional difference score was calculated, based on the weighted scores of those caregivers who did and did not have prior experience. The pooled difference score was +.13, CI +.04, +.22. There is some evidence overall for a small treatment effect on caregiver stress.

Impact on Patient Functioning

A comparison of the proportional difference scores on patient functioning for the treatment and comparison groups is shown in Table 5. There is no evidence that the Post-Hospital Support Program had any effect—positive or negative—on patient functioning. Patients in both groups showed early improvement in Activities of Daily Living (ADL), followed by a leveling off. There was very little change in the Instrumental Activities of Daily Living (IADL) scores of either group. Both groups showed some improvement in mental status in the first part of the year, with a leveling off in the 9–12 month period.

Mortality

One year after hospital discharge, 27 patients in the treatment group had died (27.5 percent), compared to 28 in the comparison group (30 percent). While there is no evidence that the "Post-hospital Support Program" decreased

TABLE 4—Comparison of Treatment and Comparison Groups on Mean Proportional Difference in Caregiver Stress Variables at Four Time Periods*

	Comparison Group		Treatment Group		Difference	95% Confidence Intervals of Difference
	a	(N)	b	(N)	a − b	
General Health Questionnaire						
0–3 months	−.23	(68)	−.32	(82)	+.09	−.06, +.24
3–6 months	−.13	(55)	−.17	(74)	+.04	−.10, +.18
6–9 months	−.25	(51)	−.28	(68)	+.03	−.14, +.20
9–12 months	−.11	(48)	−.15	(61)	+.04	−.11, +.19
Total: 0–12 months	−.30	(48)	−.35	(61)	+.05	−.14, +.24
# Caregiver Illnesses†						
0–3 months	−.22	(68)	−.25	(82)	+.03	−.09, +.14
3–6 months	−.13	(55)	−.13	(74)	+.00	−.10, +.10
6–9 months	−.09	(51)	−.21	(69)	+.13	+.04, +.22
9–12 months	−.21	(48)	−.15	(62)	−.06	−.19, +.07
Total: 0–12 months	−.26	(48)	−.20	(62)	−.06	−.20, +.08
Negative Impact of Caregiving						
0–3 months	−.19	(66)	−.31	(82)	+.12	+.03, +.19
3–6 months	−.03	(55)	+.03	(74)	−.05	−.11, +.01
6–9 months	−.06	(49)	−.03	(69)	−.03	−.12, +.06
9–12 months	+.24	(47)	+.28	(61)	−.04	−.11, +.03
Total: 0–12 months	+.00	(48)	+.04	(61)	−.04	−.12, +.04

*On all of the caregiver stress variables, a high score indicates an unfavorable outcome.
†Adjusted for previous caregiving experience.

TABLE 5—Comparison of Treatment and Comparison Groups on Mean Proportional Difference in Functional Status at Four Time Periods*

	Comparison Group		Treatment Group		Difference	95% Confidence Intervals of Difference
	a	(N)	b	(N)	a − b	
Activities of Daily Living (ADL)						
0–3 months	−.13	(66)	−.18	(82)	+.05	−.09, +.19
3–6 months	+.01	(55)	−.05	(74)	+.06	−.06, +.18
6–9 months	+.05	(50)	+.03	(69)	+.02	−.11, +.15
9–12 months	−.04	(47)	−.06	(62)	−.02	−.13, +.09
Total: 0–12 months	−.12	(48)	−.20	(62)	−.08	−.26, +.10
Instrumental Activities of Daily Living (ADL)						
0–3 months	−.02	(65)	−.05	(82)	−.07	−.23, +.09
3–6 months	+.08	(54)	.00	(73)	+.08	−.19, +.35
6–9 months	−.13	(50)	−.04	(69)	−.09	−.23, +.05
9–12 months	+.02	(47)	+.03	(62)	+.01	−.18, +.16
Total: 0–12 months	−.05	(48)	+.10	(62)	−.15	−.37, +.07
Mental Status (MSQ)						
0–3 months	+.13	(69)	+.04	(82)	+.09	−.05, +.23
3–6 months	+.25	(57)	+.20	(74)	+.05	−.11, +.21
6–9 months	+.08	(53)	+.09	(69)	−.01	−.17, +.15
9–12 months	.00	(50)	+.02	(62)	−.02	−.17, +.13
Total: 0–12 months	+.04	(50)	+.08	(62)	−.04	−.22, +.14

*On the ADL scale, a high score indicates that more assistance is needed with activities of daily living, hence, an undesirable outcome. On the IADL and MSQ indices, a higher score indicates better functioning.

mortality, there is some evidence that the program may have kept patients alive longer than those in the comparison group.

The Mantel-Haenszel odds ratios for four time periods show that, in the first nine months of the program, the odds of surviving were higher in the treatment group (OR 1.25 to 1.46). In the last three months of the 12-month follow-up period, the treatment group had slightly lower odds of surviving than did the comparison group (OR = .92). The overall OR is 1.16. These differences are small, and the confidence intervals include an odds ratio of 0 in all time periods. A Kaplan-Meier survival curve (Figure 1) shows this pattern graphically. Thus, while we cannot conclude that the treatment program reduced mortality, the data suggest that some deaths may have been postponed.

Utilization of Health Services

Hospital Days—On the basis of the hospital charts, the patients in the study used an average of 33.9 hospital days during the study. This includes the index hospital episode where the patient was picked up for the study, which averaged 18.4 days. The index hospitalization is included because the PHS (post-hospital support) program began to impact on the situation well before discharge as PHS staff worked with hospital floor staff to plan for the discharge. Thus the program may well have impacted on the length of this initial hospitalization. Table 6 shows that there was a substantial difference in the number of days patients in the treatment and comparison groups were hospitalized. Patients in the comparison group averaged 38.5 days in the hospital (of which 21 days were for the index hospitalization). Patients in the treatment group were hospitalized an average of 29.1 days (of which 16 days were for the index hospitalization). Thus the patients in the Post-Hospital Support Program had fewer hospital days, both in the initial hospitalization and in the subsequent year. Because the mortality experience of the two groups was somewhat different, these averages were also calculated, excluding all patients who did not complete 12 months in the study. These results (line 4, Table 6) serve to strengthen the conclusion.

Because confounding was a possibility in this study, the potentially confounding variables ("good" prognosis and prior caregiver experience) were correlated with hospital days. Surprisingly, prognosis was not related to hospital days. However, prior caregiver experience showed a possible relationship ($r^2 = -.04$), suggesting that patients with caregivers who had been caring for them prior to the index hospitalization had fewer hospital days. Patients of caregivers with prior experience averaged 31.5 hospital days, compared to 39.9 hospital days for new caregivers. The pooled difference score in this case was 8.2 days CI +.7, +16, suggesting that the difference was not due to confounding.

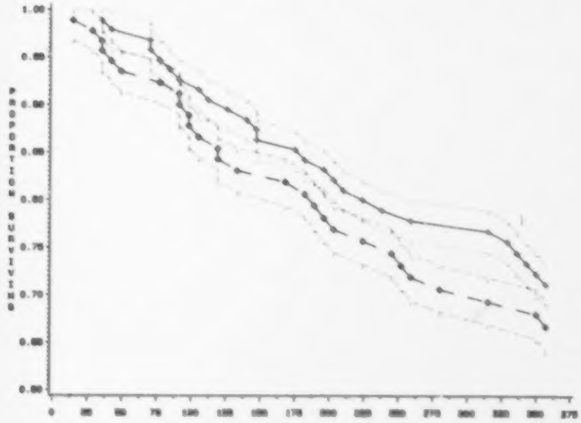

FIGURE 1—Survival Curves (with 95% confidence intervals) for Patients in the Treatment (PHS) and Comparison Groups
-✱-✱-✱- Treatment Group
⋯⋯ 95% confidence intervals for Treatment Group
◆-◆-◆ Comparison Group
-·-·- 95% confidence intervals for Comparison Group

We conclude that the patients in the Post-Hospital Support Program spent fewer days in the hospital than did patients in the comparison group. The total difference averaged about eight days. About half of this savings was due to earlier discharges following the initial hospitalization. It seems probable that hospital staff were comfortable discharging patients with high levels of need earlier, knowing that a comprehensive support program would be available. However, the savings also continued in the year following discharge. The fact that there was no meaningful difference between the groups in the average number of hospitalizations (comparison group $\bar{X} = .66$, treatment group $\bar{X} = .78$) suggests that the same factor was operating in subsequent hospitalizations. Thus the treatment program probably did not prevent hospitalizations; it shortened them.

Table 6 also shows that the patients in the treatment program used fewer nursing home days than did those in the comparison group. However, the confidence interval is too wide to conclude that this difference was not due to chance. Since the proportion of patients who entered nursing homes was about the same in the two groups (12 percent in the comparison group, 10 percent in the treatment group), the difference in nursing home use is probably due to some postponement in entry. In fact, the mean number of nursing home days for those patients in the comparison group who entered nursing homes was 146, compared to only 86 for those who entered nursing homes from the treatment group. This represents an average postponement of about two months.

Not surprisingly, the treatment patients who were in the study at 12 months had slightly more outpatient physician visits than did those in the comparison group. This probably reflects some shifting of services from inpatient to outpatient.

Cost Implications

The Post-Hospital Support Program aimed to reduce costs by reducing unnecessary health service utilization. The figures presented in Table 6 have important economic implications. In order to estimate costs under the two conditions (treatment and comparison) average costs at the Johns Hopkins Hospital (or in the Baltimore area) during the study period were gathered. At the Johns Hopkins Hospital, the average cost of a hospital day in 1985 for patients over 65 years was $1,100. It represents an average of the more expensive first part of a hospitalization and the less expensive later part. This figure includes room, board, and medical services. The average nursing home ran about $65 per day. An outpatient visit averaged $50. Applying these figures to the data for Table 6, a 6.5 day reduction in hospital days means an average saving of $7,150. Eight less nursing home days represents a saving of $520. Two more outpatient visits means an extra cost of $100 per patient. This totals to a savings of $7,670 per patient per year.

This potential saving must be balanced against the program costs. The Post-Hospital Support Program actual costs averaged $4,584.67 per patient. While the cost of similar services to the comparison group are not as readily available, they can be estimated. Using the percents receiving visiting nurse services, we estimate 1.23 home health episodes per comparison group patient. Since an average home health episode at the Johns Hopkins Home Health Agency costs $1,116, the cost of home health services can be estimated at

TABLE 6—Utilization of Health Services by Patients in the Comparison and Treatment Groups

	Comparison Group Mean (std dev.) (a)	Treatment Group Mean (std dev.) (b)	Difference (a − b)	95% Confidence Interval of the Difference
Hospital Days				
Mean days in index hospitalization	21 (16.6)	16 (14.6)	+5	+.4, +9.6
Mean days in hospital from discharge to exit from study	17 (22.6)	12 (19.8)	+5	−1.4, +11.4
Total mean days in hospital (All cases)	39 (29.9)	29 (25.5)	+10	+1.6, +18.4
Total mean days in hospital (12 months cases only)	38 (30.5)	26 (25.3)	+12	+1, +23
Nursing Home Days				
Mean number of nursing home days	17 (59.3)	9 (32.9)	+8	−.56, +21.6
Outpatient Visits				
Mean number of outpatient clinic visits	5 (15.2)	7 (12.2)	−2	−1.9, +5.9
Mean number of clinic visits (12 month cases only)	8 (20.0)	10 (14.3)	−2	−3, −1

$1,372.68 per comparison group patient. Since we cannot calculate costs for transportation (received by 23–30 percent of the comparison group patients) or respite (received by 9–14 percent of comparison group patients), we will round the figure upwards to $1,500 to roughly account for these additional costs. Using this figure, the Post-Hospital Support Program cost about $3,085 more per patient per year. If we compare this to the $7,670 saved in the lower health service utilization, we see that there remains a cost saving of $4,585 per patient per year after additional program costs are taken into account.

Discussion

The Post-Hospital Support Program for the Frail Elderly and Their Caregivers was only partially successful in meeting its goals. There was evidence of some reduction in caregiver stress, but it was not impressive. There was no evidence of improved patient functioning. There was, however, a suggestion that the program may have postponed both deaths and nursing home placements. There was also a clear pattern of evidence suggesting that hospitalizations were shortened for treatment group patients, resulting in substantial cost reductions. These went well beyond the additional program costs.

Because our design did not control the threat of history to internal validity, we cannot overlook the fact that there were substantial changes taking place in the health care system during the course of the study. Because of the introduction of DRGs (diagnostic related groups) by the Medicare Program, hospital stays were being reduced for the elderly at this time.[19] It is possible that the reduction we observed was due to the difference between the years of observation (1983–84 vs 1984–85) rather than to the effect of the treatment program. However, the average reduction in hospital length of stay for patients 65 years of age and older between 1983 and 1985 was one day.† Since, in our study, patients had an average of 1.73 hospitalizations over the study period, we would have expected a reduction of 1.73 days on the basis of history alone. Since the difference we found was almost five times this large, it is highly unlikely that the prospective pricing system could have accounted for the differences found. Also because the 1.73 days related to all patients over age 65, and the patients in our study were a very frail subset of the over 65 population, the 1.73 day figure may well be high. Even if we subtract the 1.73 expected days shorter for the treatment group, we still find at least 6.5 days difference.

Because of the quasi-experimental design and the turbulent hospital environment which existed at the time of the study, the results should be subjected to more rigorous tests before firm conclusions are drawn.

The program's lack of success in reducing caregiver stress suggests a need to reexamine both the program offered and the assumptions behind it. Low attendance at caregiver support groups may have been due to program failure (e.g., lack of skilled leadership, of adequate transportation, or respite) or to a lack of interest in this type of service by caregivers. Recent results from respite programs for caregivers show similar findings,[20–22] suggesting that perhaps the whole notion of caregiver support needs to be reexamined.

In the Post-Hospital Support Program, there were frequent conflicts between the staff and patients/caregivers. These were usually because the staff wanted the medical regimen to be followed, or wanted a higher standard of personal care to be provided, and these issues may not have been important goals for patients or caregivers. Also, when the interest of patient and caregiver conflicted, the staff tended to become advocates for the patient in spite of program goals. At staff meetings, staff frequently expressed frustration with caregivers who were not doing a "good-enough" job. It is possible that frequent visiting by this staff may have been more stress-producing than stress-reducing for caregivers. Another possibility is that the amount of support available may have been too small considering the nature of the task. One social worker told how she had spoken with a caregiver who complained that she was "so tired!" The social worker worked for several weeks to arrange help: some day care, home health aide and increased family help. When she called back to see how things were going, the caregiver again complained that she was "so tired!" Perhaps we need to examine in more depth the nature of the caregiving tasks, the interactions of patient, caregiver and agency,[23] and the psychology of mid-life women. This type of analysis may lead to the development of totally different types of "support" services.

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†Personal communication. Hyattsville, MD: Scientific and Technical Information Branch, National Center for Health Statistics, Nov. 1, 1987.

A final note on caregiving. The current study underestimated the complexity of the caregiving situation. For example, the design designated one person as a "caregiver" and one as a "patient". In fact, a number of patients had multiple caregivers—sometimes sharing responsibilities, and sometimes caring for the patient in a "serial" pattern. In some situations, it was difficult to distinguish "caregiver" and "patient" as the year progressed. When two ill spouses care for each other, people shifted in and out of these roles as their illness waxed and waned. An index, such as the Negative Impact of Caregiving, was hardly appropriate when the designated caregiver was now in the patient role. The study results also suggest that the amount of previous caregiving experience was a critical variable. Future support programs would do well to consider what type of caregiver their services are aimed at. Programs for new caregivers might need to be quite different than those for long-term providers of care.

Although this was not the program's primary focus, the reduction in health service utilization suggested by the results is intriguing. It raises the possibility that the key to reducing hospital costs lies not in what happens in the hospital itself (e.g., discharge planning) but in what happens in the community. By strengthening the caregiver, the Post-Hospital Support Program may have allowed caregivers to take on even more responsibility. This may explain why stress levels were not reduced further.

The results are promising in that they suggest that programs designed to support the frail elderly population and their caregivers in the community can be cost efficient by reducing levels of health service utilization. However, further work is needed to design programs which are also beneficial to both patients and caregivers.

ACKNOWLEDGMENTS

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APPENDIX A

Description of the Treatment Program

Treatment began in the hospital with assessments by both the Project nurse and social worker. At that point, an initial service plan would be established and caregiver education would begin. Because the initial adjustment to home care was often difficult, and patients were often very sick at this point, many patients needed frequent nursing visits in the weeks following discharge. If dressing changes were involved, up to three nursing visits a day were provided. Patients with complex medication schedules needed close monitoring and nurses helped caregivers to develop systems to simplify regimens and watched carefully to see the effects. Often they spent many hours communicating back and forth with the physicians until an appropriate medication schedule was established. If there were family problems, the social worker would become involved, helping families to develop caregiving patterns which they would manage. This often meant adding outside support. The Post-Hospital Support Program (PHS) made funds available for caregivers to hire helpers for personal care, monitoring, day care, and respite. The number of hours needed, and the supervision of outside help would be negotiated with the family, and an initial contract signed. Also, in rare cases, PHS would cover the cost of environmental modifications (e.g., bars, ramps, stair railings). Physical therapy was also provided if the staff thought it would be helpful, and it could not be covered by an outside payor.

After the initial adjustment period, each case received at least one nurse and one social work visit per month. In many cases, stable periods where only minimal assessment and monitoring were needed alternated with crisis periods, when very intensive levels of outside support were needed. The average was four nurse visits per month, and the maximum was a case which received 198 nursing visits in a year. The nurses carried a maximum of 20 cases, and the social workers 30 cases. If rehospitalization was needed, the staff would follow the patient's progress and prepare another plan for post-discharge care. Support group sessions were offered about once a month. Each session had a specific topic (e.g., "dementia", "handling caregiver stress") which opened with some prepared remarks by staff or invited speakers and ended with discussion among the caregivers. Food (usually lunch) was served, and respite and transportation arrangements were made to facilitate attendance. In spite of this, only a minority of caregivers attended those meetings (usually 10–15 caregivers). Meetings usually began with each of the caregivers introducing themselves and telling their "story". Each introduction was followed by an enthusiastic round of applause and many encouraging comments from the other caregivers (e.g., "Aren't you something!" "Your father is lucky to have you!" "You're doing everything you can.") Although not heavily attended, these groups sessions were real morale boosters for both caregivers and staff. Caregivers loved being treated "special" by the hospital. Staff enjoyed watching the caregivers interact, and they also received very heartfelt appreciation ("I don't know what I would have done without Ms. Gina [nurse]").

Throughout the duration of the program, a weekly staff meeting was held. This was attended by all staff nurses and social workers, a geriatric fellow, a psychogeriatric nurse, the researcher, and the project administrator. Each case was discussed on a periodic basis, but most of the time went to discussing problem cases, usually cases experiencing a medical crisis, or cases where the caregiving situation was deteriorating.

The program sought the best possible situation for each patient. Wherever possible, the staff tried to arrange things so that the patient could remain in the community as long as this was in the patient's best interest. In cases where the caregiving situation was poor, and did not improve in spite of repeated attempts to "prop it up" with education, added help and support, the staff would work actively to remove the patient from the home, even if it meant a nursing home placement. Although this type of situation occurred infrequently, deciding when a patient would be better off outside of the home was one of the most difficult aspects of the program for the staff.

As the end of the one year service period grew near, the staff arranged for a "discharge plan" which would provide ongoing care if needed. Those in need of skilled nursing care would be referred to Home Health Care services. Those who only needed somewhere to call if problems arose would be connected to a case-management service.

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State Funding for AIDS Patient Care Increased 160% in 1989

State government funding for AIDS patient care in fiscal year 1989 totaled nearly $65 million or 26 percent of the total non-Medicaid state funds for AIDS, according to a report released October 17 by the George Washington University AIDS Policy Center. This represents a 160 percent increase from the $25 million appropriated in fiscal year 1988. Based on information from state officials nationwide, patient care will continue to absorb a significant share of state funds as the epidemic grows and more people become infected with the human immunodeficiency virus (HIV). Twenty-six states budgeted funds for at least one type of patient care service, whereas in fiscal year 1988, only nine states budgeted funds for that purpose.

Four states (California, Florida, Massachusetts, New York), allocated more than $5 million in state-only non-Medicaid funds for patient care programs for fiscal year 1989. Eight jurisdictions (District of Columbia, Georgia, Illinois, Louisiana, Maryland, Michigan, Missouri, Pennsylvania) appropriated $1–5 million.

The high cost of Azidothymidine (AZT) has prompted at least 10 states, including such low-prevalence states as Hawaii and Utah, to set aside funds to supplement federal allocations for AZT.

The AIDS Policy Center, a component of GWU's Intergovernmental Health Policy Project, is a legislative and policy research center that provides AIDS information and related research to legislators, policymakers and the media. This report, based on the Center's national survey of state AIDS funds released August 30, 1989, is the first of several analyses of states' allocation of non-Medicaid funds for various program areas. Other reports will examine education/information, testing and counseling, research, surveillance activities, support services and administrative activities. The survey and data analyses were supported by a grant from the American Foundation for AIDS Research (AmFAR). For a copy of this report, contact the AIDS Policy Center, (202) 676-8144.

Epidemiologic Features of Treated End-Stage Renal Disease in a Large Prepaid Health Plan

Juan D. Ordonez, MD, MPH, Robert A. Hiatt, MD, PhD, and Charles P. Quesenberry, PhD

Abstract: We examined the incidence of treated end-stage renal disease (ESRD) among the two million members of the Kaiser Permanente Medical Care Program of Northern California from 1973 through 1985. The age-adjusted incidence rates rose from 33.9 to 81.7 per million person-years and showed no signs of stabilization. Rates were higher in men than in women and rose with age. Marked upward trends ($p < .001$) in the rates of treated ESRD associated with diabetic nephropathy and hypertension were noted. (*Am J Public Health* 1990; 80:47–49.)

Introduction

Since the enactment of Medicare legislation on end-stage renal disease (ESRD) in 1973, the cost to the United States government of caring for these patients has far exceeded expectations. The number of persons enrolled increased from 16,000 in 1974 to 99,000 in 1986, while the costs rose from $229 million to over $2.8 billion.[1]

Few reports document the incidence of ESRD in defined US populations, and those available from the Health Care Financing Administration (HCFA) have been limited by discontinuity in the data collection process[2] and lack of compliance with the patient history form.[1] Other regional studies encompass very short periods of time[3] or include no data on trends in incidence.[4–7]

This report describes the population treated for ESRD within the membership of the Kaiser Permanente Medical Care Program (KPMCP) of Northern California from 1973 through 1985. It is an extension of an earlier study derived from the same data sources for the years 1973 through 1977.[7]

Methods

The Kaiser Permanente Health Plan provides comprehensive hospital and inpatient services to subscribers and their dependents. The membership increased from 1.1 million in 1973 to about 1.9 million in 1986 and comprises about 25 percent of the population in the San Francisco Bay Area. The age distribution of the membership remained reasonably unchanged during the period of study. The demographic characteristics of this population in the San Francisco-Oakland Metropolitan Statistical Area are similar to those enumerated by the US census for the same area, except for some underrepresentation of persons in the extremes of the socioeconomic scale.[8] Dialysis and transplantation are generally subcontracted to outside providers.

Patient Identification

Patients were all health plan subscribers with ESRD treated with maintenance hemodialysis, continuous ambulatory peritoneal dialysis, or kidney transplantation from 1973 through 1985. Complete ascertainment was possible because patients requiring ESRD treatment were either treated at a dialysis unit in the Kaiser Permanente San Francisco facility or referred to external facilities through a central health plan office.

Medical Record Review

Medical records of all patients referred for ESRD treatment from 1972 through 1977 were reviewed in 1978,[7] those referred from 1978 through 1981 were reviewed in 1982,[9] and those referred between 1981 and 1985 were reviewed in 1986. In each study, methodology was similar. Medical records in all three periods were reviewed by a trained medical record analyst to obtain: demographic characteristics, renal diagnosis and date, whether a kidney biopsy was performed, date of referral for ESRD treatment, location of the referring facility and, when available, date of first dialysis.

Analysis

The date of first treatment was used to determine the annual age-adjusted incidence of ESRD. Only cases first treated after joining the health plan were included in the calculation of incidence rates. Rates were age-adjusted using the direct method with the 1980 US census used as standard. Time trends in the age-adjusted rates were examined using weighted least squares regression techniques.

Results

From 1973 to 1985, 1,134 patients were treated for ESRD, including 68 (6.0 percent) who were first treated before becoming health plan subscribers. Among the 1,066 diagnosed while subscribers (Table 1), the most common primary diagnoses were glomerulonephritis (25.4 percent), diabetic nephropathy (22.2 percent), primary hypertensive disease (13.4 percent), and polycystic kidney disease (7.9 percent).

The age-adjusted incidence rose in a continuous trend from 33.9 per million members in 1973 to 81.7 per million in 1985 (Figure 1). Although the most marked relative increase in incidence was among persons older than 65 years (40.0 per million in 1973 to 220.4 in 1985), the 15- to 44-year-old age group contributed the largest absolute number of patients (Figure 2).

A higher incidence in men was seen for all age groups after the first decade of life (Figure 3). The incidence in men increased steadily with age, in contrast to the incidence in women, which peaked in the 60- to 69-year-old age group.

Upward trends in the annual age-adjusted, disease-specific incidence were greatest for diabetic nephropathy ($p < .001$) and hypertension ($p < .001$). The incidence of polycystic kidney disease (PCK) increased slightly (trend, $p = .033$) and that of glomerulonephritis showed no evidence of a trend. The incidence of ESRD for all other diagnoses ranged between seven and 15 per million without any clear

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TABLE 1—Mean Age-Adjusted Incidence Rates of Treated ESRD by Diagnosis, KPMCP, 1973–85

Diagnosis	ESRD per Million Person-Years (N)		
	Men	Women	Total
Glomerulonephritis	18.3 (161)	11.0 (110)	14.6 (271)
Diabetic nephropathy	11.6 (124)	10.2 (113)	10.9 (237)
Primary hypertension	11.3 (94)	5.6 (49)	8.3 (143)
Polycystic kidney disease	4.5 (44)	3.9 (40)	4.2 (84)
Pyelonephritis	1.5 (12)	2.8 (26)	2.2 (38)
Obstructive uropathy	2.7 (20)	1.3 (16)	2.0 (36)
Collagen disease	.7 (6)	1.6 (17)	1.2 (23)
Congenital renal disease	1.2 (14)	.6 (7)	1.9 (21)
Drugs	1.2 (6)	1.0 (9)	1.1 (15)
Tumors	1.3 (11)	.4 (5)	.8 (16)
Interstitial disease	.6 (4)	.9 (9)	.7 (13)
Acute renal failure	.5 (5)	.6 (7)	.5 (12)
Other	1.7 (10)	1.6 (14)	1.6 (24)
Unknown	13.0 (99)	3.9 (34)	8.3 (133)
Total	70.0 (610)	45.6 (456)	57.2 (1066)

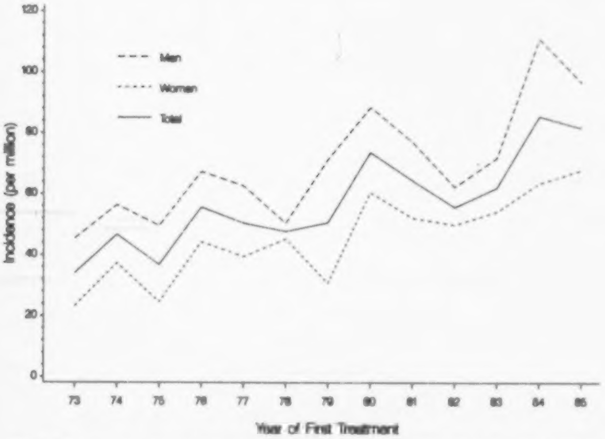

FIGURE 1—Annual Age-Adjusted Incidence of Treated ESRD in the KPMCP, 1973–85

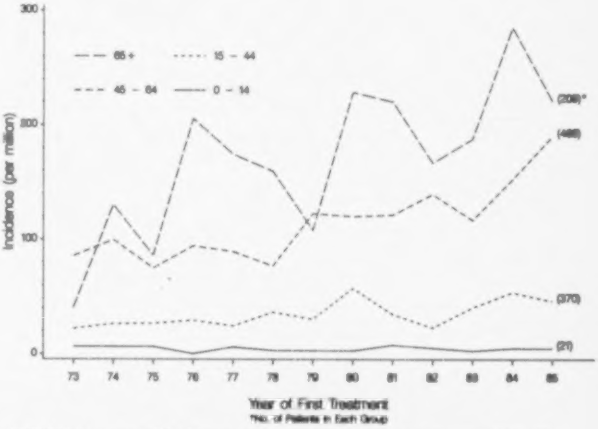

FIGURE 2—Annual Incidence of Treated ESRD by Age Group, 1973–85

trend over time.* The higher incidence of treated ESRD in the oldest age group (>65 years) was primarily associated with diagnoses of glomerulonephritis (23.4 percent) and hypertension (24.4 percent), whereas in the 45- to 64-year-old group, PCK and diabetic nephropathy comprised the largest proportion of patients (38.6 percent).

A small decrease occurred in the proportion of Blacks treated for ESRD in 1982–85 (18 percent), compared with 1973–77 (21 percent) and 1978–81 (24 percent). During the 1973–81 period, glomerulonephritis was the most frequent diagnosis among Whites and Hispanics (29 percent) and among Asians (46 percent) treated for ESRD; hypertension was the most frequent diagnosis among Blacks (26 percent). During the 1982–85 period, diabetic nephropathy became the most frequent diagnosis among all racial groups except Asians (Figure 4).

Discussion

The demographic characteristics of the patients treated for ESRD within the KPMCP in Northern California are similar to those in previous reports,[1,10] based on HCFA data. The incidence of treated ESRD in this health plan population has increased steadily from 1973 to 1985. This observation is compatible with those of Eggers, *et al*,[1] and Weller, *et al*,[11] and suggests that any conclusion that rates have stabilized[10] may be premature.

The most likely reason for the increasing incidence of treated ESRD is the continued liberalization of the criteria for the use of dialysis and transplantation by physicians, which is a consequence of improved survival and decreasing morbidity following treatment of ESRD in recent years.[12,13] Older patients with diseases more likely to be associated with systemic complications, such as diabetes and hypertension, may have benefited the most from the liberalization in criteria, because healthier patients were more likely to have been accepted for treatment in previous years. A possible increase in the incidence of insulin-dependent diabetes may also be contributing to the upward trend in incidence.[14] Finally, the improved medical management of diabetes and hypertension may have decreased competing mortality from coronary artery disease, thus increasing the number of diabetic patients reaching ESRD.

The higher incidence of treated ESRD in older age groups is compatible with the progression of chronic underlying processes that lead to terminal renal failure. The high

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*Data available on request to author.

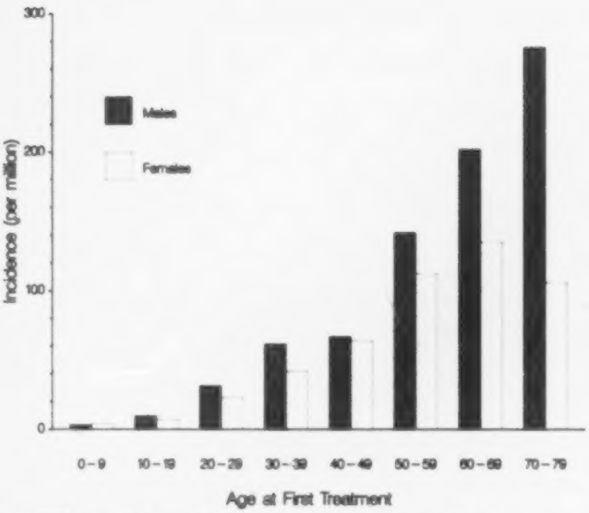

FIGURE 3—Incidence of Treated ESRD by Age and Sex, 1973–85

incidence noted in the group older than 70 years contrasts with national data from HCFA[1] and suggests that age may be less of a deterrent to use of hemodialysis among the KPMCP membership.

We were not able to calculate incidence rates by race because of the lack of data on racial composition of the health plan membership. However, the proportion of Blacks and that of Whites among those treated for ESRD may be decreasing in this Northern California population, while the proportion of Asians (and possibly of Hispanics) is increasing.[15] A higher incidence of diabetes has been reported among Hispanics, compared with other racial groups.[16]

ACKNOWLEDGMENTS

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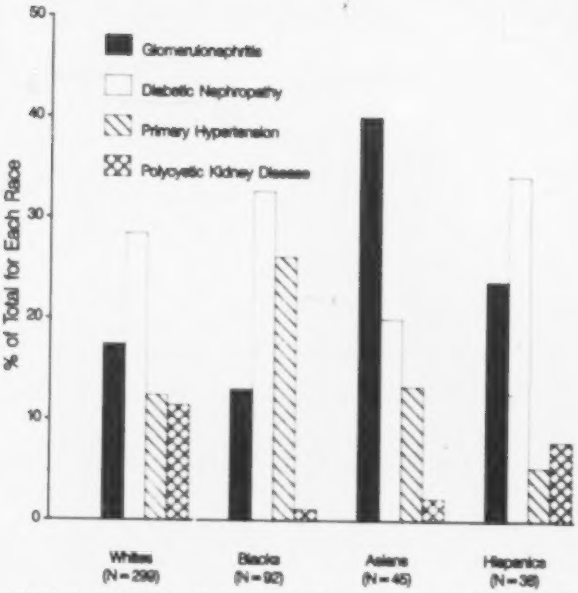

FIGURE 4—Distribution of Primary Renal Diagnosis among Treated ESRD Cases by Race, 1982–85

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An Epidemiologic Postmarketing Surveillance Study of Prescription Acne Medications

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Abstract: After the Food and Drug Administration approved a topical antibiotic for the treatment of acne, we began a postmarketing surveillance study to measure the frequency of antibiotic-associated colitis and diarrhea in acne patients treated with topical antibiotics and low doses of oral antibiotics. Pharmacists recruited 13,465 patients who presented a prescription for an acne medication, and we obtained detailed information on the use of the medications, perceived acceptability, and the occurrence of new health events from 6,453 with computer-assisted telephone interviews. Three cases (<0.1%) of antibiotic-associated diarrhea and one case (<0.1%) of antibiotic-associated colitis were confirmed. We conclude that this methodology can provide further information about the safety of a drug once it is in customary use, at a fraction of the cost of Phase III clinical trials. (*Am J Public Health* 1990; 80:50–53.)

Introduction

In November 1981, the Upjohn Company's Medication Monitoring Program^{SM} initiated a postmarketing surveillance study of prescription acne medications. Prior to the initiation of the study, Upjohn introduced a topical clindamycin phosphate product (CLEOCIN T® Topical Solution) for the treatment of acne. Many other antibiotics are available for the treatment of acne. Since pseudomembranous colitis and severe diarrhea can occur with the administration of antibiotics,[1–6] with serious or even fatal consequences, it is important to measure whether the risk of developing colitis or severe diarrhea is increased in patients with acne being treated with topical antibiotics or low doses of oral antibiotics. Information from clinical trials, informal surveys, and case reports indicates that the risk in acne patients is rather low,[7–9] but estimates of the risk under normal use are lacking. We therefore conducted a postmarketing surveillance study to estimate the frequency of colitis and diarrhea in a sample of acne patients presenting prescriptions to pharmacies.

Methods

Recruitment of Pharmacies

A pharmacy-based model of postmarketing surveillance was developed by the Medication Monitoring Program.[10] Retail pharmacies, pharmacies associated with a dermatology clinic, university student health centers, and pharmacies on military bases were recruited to participate in the study, based on willingness to participate and medication dispensing profile. Fifty-eight pharmacies participated in the study, which represented a broad geographic distribution. Pharmacists were nominally reimbursed for their participation.

To create a registry of acne medication users from which to sample for further follow-up, we asked cooperating pharmacists to explain the study to all individuals who presented a prescription for an acne medication and invite them to participate. A total of 13,465 patients were registered by pharmacists and 10,282 (76.4 percent) signed a consent form agreeing to further follow-up and access to their medical records. Participation rates within pharmacies ranged from 9.3 percent to 100 percent and did not vary markedly by type of pharmacy.

Interviews

Participants were assigned different interview forms, based upon the drug or drugs on the registry, to save cost. Patients prescribed topical antibiotics were given the complete interview, which included patterns of use, assessment of acceptability and efficacy, and the occurrence of new health events (including non-hospitalized events, such as changes in bowel habits). Participants prescribed an oral antibiotic and any other prescription acne medication were randomly assigned to the complete interview module, to a short interview which collected only patterns of use, perceived acceptability, and the occurrence of hospitalizations and pregnancies, or to no interview.

A total of 9,965 (96.9 percent of 10,282) patients was assigned to receive two computer-assisted telephone-administered interviews: the first was scheduled five weeks following enrollment and the second was scheduled to be administered eight weeks later. The interviews were scheduled to cover the interval during which antibiotic-associated diarrhea or colitis would be most likely to occur if the patient was using the drug for the first time.[11]

If contact could not be made with the patient, a certified letter was sent to his/her last known address. The individual was asked to call the study collect or return a postcard indicating whether he/she still wished to participate. Patients who did not respond to the certified card were designated lost to follow-up and a questionnaire was sent to the enrolling institution or the prescribing physician. The questionnaire was used to determine last known vital status and whether the patient had been hospitalized during the follow-up period. No deaths were reported in this study. Of the 893 (9.0 percent of 9,965) patients lost to follow-up, 15 hospitalizations for 13 (1.5 percent) patients were reported. The mean age of those lost to follow-up was 22.7 years (range: 13 to 62 years); 60.3 percent were female.

A total of 8,304 (83.3 percent) patients completed the first interview (any module) (Table 1). Only data from the 6,453 patients who completed the full set of questions will be presented in this report. Patients were not paid for their participation.

The mean age of the participants was 22.9 years (range: <1 year to 80.6 years); 63.7 percent were female. The mean age of those not signing a consent form authorizing further follow-up was 23.4 years (range: 2.9 years to 78.1 years); 50.7 percent were female. Of all females registered, 80.2 percent

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TABLE 1—Interview Status of Study Participants

	Frequency	% of Total
Signed Consent	10282	(100.0)
Assigned to interview	9965	(96.9)
First interview completed	8304	(80.8)
First interview not completed	1661	(16.2)
Participation terminated to reduce interviewer backlog	685	(6.7)
Lost-to-followup	521	(5.1)
Registry received later than protocol requirements	199	(1.9)
No telephone	71	(0.7)
Withdrew consent	58	(0.6)
Did not use acne medication	48	(0.5)
Moved outside United States	34	(0.3)
Hostile/abusive respondent	26	(0.3)
Data processing error	19	(0.2)
Second interview completed	7424	(72.2)
Second interview not completed	880	(8.6)
Lost-to-followup	372	(3.6)
Data processing error	247	(2.4)
Moved outside United States	129	(1.3)
Withdrew consent	104	(1.0)
Hostile/abusive respondent	28	(0.3)

NOTE: This table includes only those registries in which patients agreed to participate in the study (76.4% of total registries).

agreed to participate compared to 72.8 percent of the registered males.

Investigation of Health Events

Hospitalizations, pregnancy outcome, and diarrhea requiring a physician contact were investigated further. A letter of inquiry was sent to the physician or institution requesting confirmation of dates, reasons for contact or hospitalization, and the final diagnosis. If insufficient information was returned, the report was referred to field investigation for resolution. The determination of the appropriateness of each investigated diagnosis was conducted by the medical monitor in a blinded fashion.

The criterion to define episodes of diarrhea was the patient report. No attempt was made to characterize the frequency or consistency of the stool. Clinically significant diarrhea was defined as that requiring contact with a physician and was investigated through the patient's physician. Patients who reported having the diarrhea at the time of the interview were contacted two weeks later and queried about the episode.

Results

Acne Medication Use

Multiple drug use was common, with individuals reporting an average of 2.6 acne drugs during the follow-up period. Of the 6,453 patients, 5,832 (90.4 percent) reported using at least one antibiotic during the follow-up period: 1,361 (23.3 percent) reported using only an oral antibiotic; 2,021 (34.7 percent) reported using only a topical antibiotic; and 2,450 (42.0 percent) reported using both an oral and a topical antibiotic.

Reports of Diarrhea

A total of 276 (4.3 percent) patients reported the occurrence of an episode of diarrhea during the follow-up period. The mean age of those reporting diarrhea was 23.8 years (range: 13.2 years to 53.4 years); 73.9 percent were female. A total of 4.8 percent of the female patients reported an occurrence of diarrhea compared to 3.3 percent of the male patients.

Of these 276 patients, 252 (91.3 percent) reported exposure to one or more antibiotics during the follow-up period. Table 2 presents the frequency of reported diarrhea by antibiotic exposure. Due to the large number of different types of antibiotics prescribed and the relatively small number of diarrhea cases, the exposure categories were collapsed to topical antibiotics and oral antibiotics. The rate of reported

TABLE 2—Frequency and Characteristics of Episodes of Diarrhea Reported by Acne Study Patients by Antibiotic Exposure

	Oral and Topical Antibiotics	Topical Antibiotics Only	Oral Antibiotics Only	No Antibiotic Exposure	Total
Total number of patients exposed	2450	2021	1361	621	6453
Number reporting diarrhea (%)	110 (4.5)	98 (4.8)	44 (3.2)	24 (3.9)	276 (4.3)
Associated Symptoms* (%)					
Blood/Mucus in stools	17 (15.5)	19 (19.4)	6 (13.6)	3 (12.5)	45 (16.3)
Abdominal cramps	63 (57.3)	68 (69.4)	28 (63.6)	14 (58.3)	173 (62.7)
Nausea	59 (53.6)	54 (55.1)	23 (52.3)	14 (58.3)	150 (54.3)
Fever	32 (29.1)	28 (28.6)	9 (20.5)	7 (29.2)	76 (27.5)
Associated Factors*					
Shared quarters with someone with diarrhea	26 (23.6)	30 (30.6)	4 (9.1)	3 (12.5)	63 (22.8)
Changed diet prior to onset	29 (26.4)	29 (29.6)	18 (40.9)	7 (29.2)	83 (30.1)
Trip prior to onset	23 (20.9)	29 (29.6)	11 (25.0)	4 (16.7)	67 (24.3)
Similar diarrhea previously	65 (59.1)	68 (69.4)	22 (50.0)	14 (58.3)	169 (61.2)

*Patient may report multiple symptoms/factors
( ) % of number reporting diarrhea

diarrhea was slightly higher among users of topical antibiotic preparations than among users of oral antibiotics (RR = 1.5; 95% CI = 1.06, 2.13).

Further information was obtained about each reported episode. Table 2 shows the distribution of reported symptoms associated with the diarrhea by drug exposure. Blood and mucus in the stools were not asked separately. There were no significant differences in the frequency of reported symptoms associated with the diarrhea episode by antibiotic exposure. Patients were also asked whether they ever had diarrhea like this before, whether any individuals with whom they share living quarters had diarrhea around the same time, whether there had been any significant changes in diet prior to onset, and whether they had taken any trips prior to onset. The frequency of diarrhea reported by the season of the year did not vary significantly.

Twelve (0.2 percent) of 6,453 patients reported discontinuing a drug due to diarrhea; four patients discontinued oral tetracycline, two patients each discontinued topical clindamycin phosphate, oral erythromycin, and oral minocycline, and one patient each discontinued topical meclocycline and topical clindamycin HCL. Of the 3,811 patients exposed to at least one oral antibiotic, eight (0.2 percent) reported discontinuing an oral antibiotic due to diarrhea. Of the 4,471 patients exposed to at least one topical antibiotic, four (<0.1 percent) reported discontinuing a topical antibiotic due to diarrhea.

Eighty-two (1.3 percent) of 6,453 patients reported 85 contacts (visits or telephone contact) with a physician about the diarrhea episode. Confirmation of a diagnosis of diarrhea was obtained for 46 (54.1 percent) of the contacts. Three physician-confirmed cases of antibiotic-associated diarrhea included one case associated with tetracycline (oral), one case associated with erythromycin (oral), and one case associated with minocycline (oral). The causal association was based solely on the medical monitor's review of the physician's diagnosis obtained from the patient's records. No further confirmation of the diagnosis was obtained.

Colitis

Of nine (0.1 percent) patients reporting colitis, four of them reported no contact with a physician. Follow-up of the remaining five cases revealed only two cases of confirmed colitis, neither drug-related. Two other patients reported diarrhea, and follow-up with the treating physicians revealed oral tetracycline-associated colitis and stress-related colitis. Neither patient was hospitalized for the condition.

Hospitalizations

A total of 92 (1.4 percent) patients reported 93 hospitalizations during the follow-up period. The mean age of those hospitalized was 23.9 years (range: 11.1 years to 60.8 years); 63 (68.5 percent) were female. Table 3 presents a summary of the discharge diagnoses for 75 of the hospitalizations. Eighteen (19.4 percent) of the reported hospitalizations could not be confirmed. Two patients reported being hospitalized for diarrhea. One of the reports could not be confirmed and the discharge diagnosis for the other hospitalization was acute viral gastroenteritis.

Pregnancy

A total of 41 (1.0 percent) women reported being pregnant during the follow-up period. Their mean age was 24.8 years (range: 17.5 years to 34.3 years). Twenty-six (63.4 percent) of the deliveries were confirmed and four women volunteered that they had an abortion. Of the 26 confirmed births, 20 infants were normal, one infant was premature, one infant had hyperbilirubinemia, and the status of the remaining four cases was unknown. In all confirmed cases, the women reported discontinuing their acne medications during the pregnancy.

TABLE 3—Distribution of Confirmed Discharge Diagnoses for Reported Hospitalizations among Acne Study Group

Discharge Diagnosis*	Frequency	Column %
Musculoskeletal Disorders	21	18.4
Infections	18	15.8
Trauma	17	14.9
Gastrointestinal Disorders	17	14.9
Gynecologic Disorders	12	10.5
Respiratory Disorders	5	4.4
Neuropsychiatric Disorders	5	4.4
Urinary Tract Disorders	4	3.5
Other, NEC	15	13.2
Total	114	100.0

*Diagnoses categorized with SNOMED[14]

Discussion

This approach relies on the pharmacist or other pharmacy personnel to recruit patients. Probably not all patients who presented a prescription for an acne medication were invited to participate in the study; possibly invitation to participate was not random. For example, pharmacy staff may have invited patients only when they were not hurried or busy, or invited only patients who appeared friendly. We oriented the pharmacy personnel in person and frequently monitored their progress for significant variations. We did not audit the prescription records within each pharmacy, although a study in progress will do so. The impact of incomplete ascertainment or selection of only particular types of patients is unknown, but if the likelihood of invitation was not related to the medication prescribed, the comparisons of side effects among medications would not be affected.

We found that the patients used many other acne drugs that did not appear on the study registry. Since we designed the interviews to collect extensive information only about the registry drugs, we lacked important data about first time vs repeat use, length of prior use, dose and regimen.

The accurate determination of the presence of diarrhea is critical to the interpretation of the results. Stool collections or observations to accurately characterize the presence of diarrhea are generally not achievable in a general population-based study.[12] Daily diaries were thought not to be feasible due to the nature and the setting of the patient recruitment. We were left with the patient recalling the occurrence of diarrhea over a five- to eight-week period of time. We felt that clinically significant diarrhea would result in a contact with a physician and thus be available to the study by confirmation through medical records. Diarrhea serious enough to prompt the patient to seek a physician was confirmed in 46 (0.7 percent of 6,453) patients.

Despite some limitations imposed by the study design, our results are not substantially different from published data. Stergachis reported four cases of antibiotic-associated colitis that required hospitalization for treatment in a population of approximately 280,000 individuals.[13] We identified no cases of antibiotic-associated colitis that required hospitalization for treatment and one case of antibiotic-associated colitis that did not require hospitalization.

We conclude that while the occurrence of antibiotic-associated colitis and diarrhea appears to be rare in this patient population, appropriate education of the patient by the physician and the pharmacist to the potential side effects of treatment appears to be warranted.

Although epidemiologic postmarketing studies have been recommended to get further information about the safety of a drug once it is in customary use, few have been completed. We found pharmacist and patient cooperation to be high. Of the 10,282 patients who agreed to participate, we obtained complete data from 6,453 (62.8 percent). We initially assigned participants to long or short interviews to minimize the costs, but found the difference in the costs of obtaining complete information compared to partial information was inconsequential, once the patient was enrolled.

This methodology requires a sophisticated organization made up of epidemiologists, physicians, computer programmers, biostatisticians, and field personnel. Once the organization is in place, the cost of conducting a study such as this is very reasonable. Borden and Lee estimate that the cost of a study with this design is about 1 percent of that of the standard clinical trial.[10]

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AAAS Meeting to Feature Wide Range of Science, Medicine and Technology

A broad look at science—from the developing brain to exploring Mars—will be the focus of the upcoming 156th annual meeting of the American Association for the Advancement of Science (AAAS) when it convenes in New Orleans, Feb. 15–20, 1990.

Over 250 symposia will showcase what is happening in science and how that affects everyone. Among the topics featured will be neurotransplants, oil contamination in Prudhoe Bay, reducing the toll of natural hazards in the 1990s, exploration of Mars, progress in AIDS treatment, psychology of jury decision-making, climate change, chaos, effects of substance abuse on women, misconduct in science, physiology of exercise, computer viruses, and emotions and the developing brain.

A special feature of the meeting will be the inclusion of two separate three-day seminars on protein folding and the biology of parasitism; these symposia require an additional registration fee. Another feature of the meeting—"Public Science Day"—takes place on the opening day, February 15. The event will include workshops, demonstrations, and conversations with scientists. Public Science Day activities will feature programs to involve local students and their families in math and science. Also on Thursday, February 15, several thousand New Orleans area students will participate in a day-long "Science Encounters" program of lectures, demonstrations, and opportunities for young people to meet and talk with individual scientists and engineers. The AAAS Science and Technology Exhibition, February 16-19, will be at nearby Rivergate Exhibition Center. Poster sessions will also be featured adjacent to the exhibits.

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Dependency on Soup Kitchens in Urban Areas of New York State

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Abstract: We studied the dependency of persons on soup kitchens in Albany, Buffalo, Rochester, Syracuse, and Westchester County, New York. Seventeen percent of the meal recipients were homeless, 62 percent lived in apartments or houses, 20 percent were working, 40 percent were women, and 17 percent had a child in their household. Fifty-nine percent started eating at the soup kitchen more than a year ago, and 51 percent ate five or more meals at soup kitchens in the last week. Most reported they came to the soup kitchen because of economic problems or lack of food; 93 percent had incomes below the poverty threshold. Most used some government food program; 48 percent received food stamps. Utilization of soup kitchens and other programs differed between men and women and between households with and without children. (*Am J Public Health* 1990; 80:57–60.)

Introduction

The number of meals served by a national sample of private emergency food programs increased by 21 percent from 1983 to 1984 and by 17 percent from 1984 to 1985.[1] In New York State, other than New York City, the estimated number of meals served monthly at soup kitchens was 204,334 in 1987;* this represents an increase from estimates for 1985 and 1986.[2] Other reports have documented the increased use of emergency food programs, discussed possible causes for it, and recommended policies to address the food needs of the users of emergency food programs.[3–6] However, there is little information about the people who are dependent on these programs.

This paper seeks to answer the following questions:

- What were the characteristics of persons dependent on soup kitchens in five urban areas of New York State outside of New York City?
- What was the extent of and reasons for their dependency on soup kitchens?
- What other food programs did they use?
- What were their perceived needs for health and other services?
- Were there differences in the patterns of dependency between men and women and between meal recipients who lived in households with and without a child?

Methods

Selection of Sample

In 1987 a survey was conducted of soup kitchens that served meals to the public during April in Albany, Buffalo, and Westchester County, and during October in Rochester and Syracuse. Of the 34 soup kitchens in these urban areas, 28 participated in the study.

It was not possible to take a random selection of guests from the sites because the total number of unduplicated guests who ate meals at the sites was not available. However, since the total number of meals served by the sites was available, meals were sampled, and the soup kitchen guests served those meals were interviewed. Because meals were sampled, guests who ate more meals at soup kitchens than others had a greater likelihood of being interviewed.

The goal was to select 20 meals at each site. An average of 17.9 meals were selected per site. Because an earlier study[7] found that soup kitchens served a greater number of meals at the end of the month, more meals (70 percent) were selected from the last half of the month.

A variety of mealtimes for interviewing guests was selected, and at each selected mealtime, the recipients of two or three randomly selected meals were interviewed. If the recipient was a child, the accompanying adult was interviewed. If a recipient refused to be interviewed, or if someone in his or her household had already been interviewed, the person next to that individual or the recipient of the next randomly selected meal was interviewed. Of the 501 soup kitchen guests interviewed, 87 percent were the recipients of the original randomly chosen meals.

Data Collection and Analysis

The guests were asked a series of primarily closed-ended questions about their demographic characteristics, household composition, sources and amounts of household income, use of health services, and use of other food programs. Each interview took about 25 minutes.

Unweighted frequency tables were created from the sample data and the chi-square statistic applied.[8] For the frequencies and means of the distribution of single variables, combined urban estimates were calculated from the sample data by combining site estimates so as to weight for the number of meals served in each site and in each urban area, assuming that the guests at nonparticipating soup kitchens were similar to those at participating soup kitchens. The combined urban estimates represent the monthly total of soup kitchen meals (82,449) served in the five urban areas.

To estimate variances and standard errors, and to study the implications of varying sampling designs, the betabinomial distribution[9,10] was used to model percentage estimates. The method of moments[9] was applied to 37 dichotomous variables and adjusted for the finite population of sites.[11] Analysis of variance was used to determine what proportion of the among-site variance was due to within-city and among-city variances.

To determine to what degree the results based on a sample of meals could be generalized to the population of soup kitchen guests, estimates were recalculated for those variables related to the probability of a guest being selected into the sample, i.e., each response was multiplied by the

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reciprocal of the number of times each guest had eaten at soup kitchens in the last week. Responses for those who ate more than 10 meals were weighted by 1/10 so as not to underweight the responses of frequent users.

Results

For 34 of the 37 variables analyzed using the betabinomial distribution, more than half of the total variance was due to the among-site variance. The standard errors estimated from the percentages in the total sample ranged from 1.1 percent to 2.9 percent. Therefore, 3 percent can be considered as a general (conservative) estimate of the precision of the percentages reported in this paper. The variation within cities was found to be greater than the variation among cities. For 26 of the 37 variables examined, the among-city variance was zero. For the remaining 11 variables, the ratios of the among-city variance to the within-city variance ranged from .07 to .45, indicating that the among-city variance was at most less than one-half the size of the within-city variance.

When the estimates based on meals were compared to the new estimates based on people, only five variables had estimates that differed by 5 percent or more. The estimates based on meals were larger than those based on people for eating at another soup kitchen (21 percent vs 12 percent) and homelessness (17 percent vs 12 percent), and smaller for women (40 percent vs 45 percent), children in the household (17 percent vs 22 percent), and another adult in the household (12 percent vs 25 percent). Therefore, except for these few variables, the reported estimates for the population of soup kitchen meals can also be interpreted as estimates for the population of soup kitchen guests. Throughout the rest of the paper, the results based on the sample of meals are reported for meal recipients.

- *Who was dependent on soup kitchens?*

Demographic Characteristics

The majority of the meal recipients were men (60 percent). The average age of meal recipients was 43.6 years and 20 percent of them were 60 years old or older. Most of them were White (65 percent) although 29 percent were Black. Their educational attainment was generally low; 52 percent did not have a high school diploma. Most of the meal recipients lived alone or in a shelter or group home (70 percent).

Employment

Only 20 percent of the meal recipients were employed. Although 37 percent of the unemployed were looking for work, many of the unemployed were disabled (32 percent), retired (19 percent), raising a family (11 percent), or attending school or a training program (5 percent). Meal recipients who were employed earned on average $5.47 per hour and held labor (45 percent) and service (30 percent) positions (Table 1).

Income

Ninety-three percent of the meal recipients' household incomes in the previous month were less than 100 percent of the poverty level and 55 percent were less than 50 percent of the poverty level (Table 2). The most frequently reported sources of household income were public assistance other than Aid to Families with Dependent Children (AFDC) and work relief (28 percent), Supplementary Security Income (SSI) (22 percent), Social Security (22 percent), and wages (21 percent).

TABLE 1—Types of Employment of Soup Kitchen Meal Recipients[*]

Type of Employment	Examples	Percentage of Meals for Those Employed
Laborer	Day labor, collecting bottles and cans, odd jobs	45
Service	Household worker, wait person, janitor	30
Operative	Delivery person, cab driver, assembly worker	16
Clerical	Teacher's aide, clerk, secretary	7
Crafts person	Carpenter, mechanic	7
Professional/Manager	Substitute teacher	3

[*] These percentages do not add to 100 because more than one response was possible. Ninety-six sampled meal recipients responded to this question.

TABLE 2—Household Income of Soup Kitchen Meal Recipients in Previous Month

	Percentage of Meals
Monthly Household Income (does not include food stamps)	
$0–150	27
$151–300	22
$301–600	41
$601–$1000	7
$1001 +	4
Average (mean) per person household income $275	
Sources of Household Income	
Other public assistance	28
Social Security	22
SSI	22
Wages	21
AFDC	9
Work Relief	9
Retirement	4
Veterans' disability	2
Workers' compensation	2
Unemployment	2
Child support	1

[*] These percentages do not add to 100 because households received income from more than one source.

Housing

The problems that bring people to soup kitchens affect a broader segment of the population than just the homeless. Sixty-two percent of the meal recipients lived in apartments or houses. Only 17 percent were homeless—persons who had lived in a shelter (11 percent), or on the streets or in no one place (6 percent) during the previous month,[**] and 20 percent lived in a room or group home. However, other indications of housing instability were that 15 percent of the meal recipients had moved since the month before, and 7 percent were living with unrelated people.

Thirty-one percent of the meal recipients who lived in apartments, houses, or rooms had housing that was partially subsidized or paid for by the government. It was not clear what portion of the rent of guests with subsidized housing was subsidized. However, for those without subsidized housing, the average monthly household rent was $193 and comprised 57 percent of the average household income.

[**] This percentage would have been greater if soup kitchen sites that served meals only to residents of their shelters had been included in the sample.

- *What was the extent of and reasons for dependency on soup kitchens?*

Soup kitchen meal recipients tended to use soup kitchens often and over a long period of time. Fifty-one percent of the meal recipients had eaten five or more meals at soup kitchens during the previous week, and 59 percent had started eating at the soup kitchen more than a year ago.

The most common reasons that meal recipients gave for eating at the soup kitchen were the need for food (71 percent), including such reasons as no place to cook and nowhere else to eat, and economic problems (48 percent), such as no money (Table 3). Eighty-nine percent gave at least one of these reasons. However, 41 percent of the meal recipients also gave social reasons such as the need to be among people and the support of the soup kitchen staff.

- *What other food programs were used by soup kitchen meal recipients?*

Public Programs

Government food programs were used by almost two-thirds (64 percent) of the meal recipients' households, and many of these households received food from more than one program. Forty-eight percent of the soup kitchen meal recipients' households received food stamps in the previous month, 54 percent of the meal recipients whose households did not receive food stamps had received food stamps sometime in the past, 8 percent had applied for food stamps but did not receive them, and 38 percent had never applied for them.

Meal recipients' households that received food stamps received on the average $80 in food stamps in the previous month, and 96 percent of them had household incomes that were below poverty level even when the value of their food stamps was included as income.

Forty-two percent of the meal recipients' households had received food from a free food, cheese, or other government surplus distribution in the previous month. Of those meal recipients' households who may have been eligible for the program, 29 percent received food from a senior center, 57 percent received food from a free or reduced price school breakfast, and 66 percent from a free or reduced price school lunch program, and 37 percent received the Special Supplemental Food Program for Women, Infants, and Children (WIC) benefits.

TABLE 3—Meal Recipients' Reasons for Using Soup Kitchens*

Reasons	Example	Percentage of Meals
Food	No food, no place to cook, don't know how to cook, special diet, nowhere else to eat/go	71
Economic reasons	No job, waiting for public assistance, no money	48
Social reasons	Love the people here, get out, lonely	41
Program	Shelter, religious services, referred to program, to get help, volunteer	12
Convenience	Close by, kill time, break monotony, other site closed, save time on cooking	11
Miscellaneous	Family difficulty, medical problems, for exercise, old	4

*These percentages do not add to 100 because meal recipients gave more than one reason.

Private Programs

Many meal recipients had received food from other private food sources in the previous month: food pantries (33 percent), shelters that served meals only to residents (9 percent), and credit from a grocery store (3 percent).

- *What were the meal recipients' perceived need for other services?*

Housing

Many of the meal recipients lived in housing that lacked basic facilities. Sixty-one percent of the meal recipients did not have access to a phone, 29 percent did not have access to a working refrigerator, and 28 percent did not have access to a working stove or range.

Health Services

The most frequently mentioned forms of health insurance were Medicaid (56 percent) and Medicare (21 percent). Twenty-three percent of the meal recipients had no health insurance, and 19 percent reported their families had medical expenses in the previous year that were not covered by insurance. More of the meal recipients reported a need for dental services that their family could not afford (18 percent) than a need for medical services (6 percent).

Access to Food

Fifteen percent of the meal recipients had difficulty getting to a grocery store; 11 percent had problems with transportation, and 7 percent had health or mobility problems that made it difficult for them to get to a store.

- *What differences were there in the dependency of subgroups of guests?*

Reports that have noted an increase in soup kitchen use have also noted an increase in the number of families with children eating meals at soup kitchens.[1,3,4] In the cities included in this study, 40 percent of the meal recipients were women and 17 percent of the meal recipients' households included children.

Men vs Women

The men and women eating meals at soup kitchens differed in that women were more likely to have health insurance (86 vs 75 percent), to have received AFDC (17 vs 1 percent), or SSI (31 vs 20 percent), and to have received food stamps (64 vs 39 percent), surplus foods (58 vs 38 percent), and food from a food pantry (44 vs 32 percent). Men had an average per person household income of $328 compared to $255 for women. Men were more likely to have received food from a shelter in the previous month (15 vs 6 percent), and ate more frequently at soup kitchens in the previous week (4.7 vs 3.2 meals).

Guests Living with and without Children

Soup kitchen meal recipients who lived with children were on average younger than meal recipients who did not live with children (35.7 vs 46.5 years), had lower average per person household incomes ($146 vs $330), were less likely to receive Social Security (12 vs 29 percent) and were more likely to receive AFDC (37 vs 1 percent), other public assistance (36 vs 25 percent) and food stamps (74 vs 43 percent), and were less likely to eat at shelters (2 vs 14 percent). They had on average eaten fewer meals at the soup kitchen in the last week (2.8 vs 4.4) and had been eating at the soup kitchen for a shorter period of time (16.7 vs 23.0 months).

Discussion

For future studies of soup kitchen meal recipients, the large among-site variance suggests that sampling from more sites could reduce the variance of estimates. For example, given a population with 100 sites, the standard error of an estimated percentage based on a sample of 40 meal recipients from five sites would be about 11 percent compared to only 6 percent for a sample of 10 meal recipients from 20 sites.

A variety of people other than single, homeless, males were found to eat at soup kitchens. Subgroups differed in their use of soup kitchens, their use of other programs, and their resources. Men and guests who do not have children in their households may be less likely than others to be, or perceive themselves to be, eligible for government benefits and other programs or more likely to prefer to eat at soup kitchens. Whether the goal is to reduce the number of people having to use soup kitchens or to serve the unmet needs of soup kitchen guests, differences in the utilization of programs among these subgroups should be kept in mind.

We also found that a large number of soup kitchen meal recipients ate often at soup kitchens and had been eating at them for a long period of time. To the degree that guests do not use soup kitchens as a source of "emergency food" but eat at them often and over a long period of time, more attention needs to be given to the nutritional quality of soup kitchen meals.[12]

Over one-half of the meal recipients had no access to a telephone and 18 percent stated their families needed dental services which they could not afford. The use of a telephone is an important resource for obtaining employment as well as interacting with social services to apply for and receive benefits. Neglected dental problems may interfere with soup kitchen guests' abilities to eat adequate diets.

Many soup kitchen meal recipients were participating in government food programs, such as food stamps, and still eating at soup kitchens. This suggests that government food programs may not provide all persons with sufficient resources to obtain enough food. Most of those receiving food stamps reported incomes below the poverty threshold even when their food stamp benefits were considered as income. Many of the meal recipients gave a social reason as one of their reasons for coming to the soup kitchens. Participation in a government food program, such as food stamps, may not substitute for the feeling of belonging and the social support that many soup kitchen guests receive at soup kitchens.

More research is needed on soup kitchen meal recipients to determine: why more than one-half of them were not using food stamps; how housing problems other than homelessness, such as the cost of housing and the lack of basic facilities, e.g., a refrigerator, limit their access to food; and what other services are needed besides meals to improve their situations.

The findings from this study of soup kitchen guests indicate that any policies or programs to address the needs of soup kitchen guests must look past oversimplified views of soup kitchen guests, e.g., the homeless, singles, males, persons who have not accessed public programs, or persons who are eating at soup kitchens because of a temporary emergency. Soup kitchen guests include a variety of persons with a variety of patterns of and reasons for dependency on these programs.

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Predictors of the Initiation of Psychotherapeutic Medicine Use

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Abstract: This study examined the hypothesized predictive association between illicit drug use and the onset of prescribed psychotherapeutic medicine use. The effect of psychiatric symptoms on this relationship was also explored.

Data were gathered retrospectively, through standardized household interviews conducted in 1981 for the Baltimore site of the National Institute of Mental Health Epidemiologic Catchment Area Program. For this analysis, the sample was restricted to 1,716 respondents ages 18–44 at the time of interview, drawn by probability sampling area household residents of all adult ages. Using Cox proportional hazards models, the onset of psychotherapeutic medicine use was modeled as a function of illicit drug use and psychiatric symptoms, both defined as time-dependent covariates. Newly observed in this analysis was an independent association of illicit drug use with psychotherapeutic medicine use, which was unexplained by psychiatric symptoms (relative hazard = 2.61, 95% confidence interval (CI) = 2.15, 3.18). Adjustment for age of onset of alcohol intoxication did not alter the impression that illicit drug users were more likely to initiate use of prescribed psychotherapeutic medicines (relative hazard = 2.52, 95% CI = 2.02, 3.11). (*Am J Public Health* 1990;80:61–65.)

Introduction

Studies of drug sequencing have focused on detecting patterns in the age of onset of drug and medicine use among adolescents and young adults.[1–3] This research has identified a modal sequence of onset beginning with alcohol and cigarette use, followed by marijuana use, and then use of other illicit drugs. In one of the latest studies on this issue,[4] initiation of marijuana use was found to precede initiation of prescribed psychotherapeutic medicine use.

Based on these results, researchers hypothesized[5,6] that there might be a relationship between illicit drug use and psychotherapeutic medicine use. Yamaguchi and Kandel reported that illicit drug use increased the risk of initiating psychotherapeutic medicine use, using retrospective reports of the age of onset of different types of drug use. They then showed that the presence of depressive symptoms reduced this relation between the substances among women but not among men. Other studies lending support to the potential for symptoms to act as an explanatory link are those which found that relief of symptoms was a major reason given for use of illicit drugs [7,8] and those that show a direct relation between psychotherapeutic medicine use and psychiatric distress measured cross-sectionally.[9–12]

In addition to the role that psychiatric symptoms or distress may play in influencing the use of illicit drugs and psychotherapeutic medicines, other possible etiologic mechanisms have been suggested. Raveis and Kandel [13] proposed that a change in substances used over the lifespan may be part of a maturation process that coincides with the assumption of adult roles, such as establishing a residence, beginning a job, and marriage. In a study of young adults during the three years after high school graduation, Bachman, *et al*,[14] found that illicit drug use decreased from high school levels for those who were married, increased for those who were single and living away from home, and was unchanged for those who remained with their parents.

Drug use also may be understood as a learned behavior, in which reinforcing consequences of illicit drug use later influence the taking of prescribed psychotherapeutic medication.[15] Further, in a study by Newcomb, *et al*,[16] illicit drug use was correlated with deviance and with a sensation-seeking trait in adolescents, both of which could account for a relation between illicit drug use and prescribed psychotherapeutic medicine use.

In this study we hypothesized that if the data were to support such a relation, this would be due to the effect of psychiatric symptoms, such as anxiety or depression, on both forms of drug use. We also explored the effect of the sequence of onset of illicit drug use and psychiatric symptoms upon the initiation of psychotherapeutic medication. This allowed us to compare individuals who experienced symptoms prior to illicit drug use with individuals who used illicit drugs and then developed symptoms, in terms of their likelihood of subsequent psychotherapeutic medicine use. Sociodemographic variables that may influence the occurrence of psychotherapeutic medicine use were treated as potentially confounding or modifying covariates.

Our study expands upon prior research in this area by its use of a more heterogeneous sample with a wider age range, and by broadening coverage of psychiatric symptom categories as opposed to earlier research in which only one symptom category, such as depressive symptoms, was measured. Furthermore, incorporation of age of onset data into the analyses, for the first time allows us to examine the effect of sequence of onset of symptoms, illicit drug use, and prescribed psychotherapeutic medicine use.

Methods

Study Sample

The study sample consisted of subjects selected for the Baltimore site of the National Institute of Mental Health Epidemiologic Catchment Area Program. The details of the survey design have been reported elsewhere.[17] Briefly, the population from which the sample was drawn was comprised of adult household residents of the Eastern Baltimore Epidemiologic Catchment Area (ECA), approximately the eastern one-third of Baltimore City. Respondents were chosen for inclusion in the sample using area probability sampling. Household residents eligible for selection were all persons ages 18 or over at the time of contact in 1981, with the ECA defined as their usual place of residence. The survey completion rate for designated respondents ages 18–44 ranged

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from 88–93 percent for race other than White and 80–89 percent for Whites.[17]

Of the 3,481 people interviewed, all subjects ages 18–44 at the time of interview were included in the present analysis (n = 1,716). The upper age limit for inclusion in the study was 44 years, since onset of illicit drug use after this age was quite rare.[18] The sample consisted of 1,048 females (61 percent) and 668 males, with 916 Whites (53 percent), 746 Blacks (44 percent), and 54 (3 percent) of other races. Seventy-six percent of the sample was in the married or single categories at the time of interview, and 45 percent had annual household incomes under $15,000. Twenty-seven percent of the respondents reported that they had sought help for a nervous or emotional problem at least once in their lifetime.

Study Variables

Information on the study variables was obtained from standardized interview questions covering demographics, health services utilization, the Diagnostic Interview Schedule, and a detailed survey of medicine use. Detailed information on the survey interview has been published in a monograph.[17]

Psychotherapeutic medicines were defined as those used primarily to treat, control, or prevent mental and behavioral disturbances, or to alleviate symptoms of psychopathology. Included in this definition were prescribed anti-psychotic agents, anti-anxiety agents, anti-depressants, stimulants, daytime sedatives, and hypnotics. Excluded were medicines with some psychotherapeutic action that are primarily prescribed for other purposes, such as analgesics, and over-the-counter medicines. Respondents were considered to be positive for psychotherapeutic medicine use if they stated that they had taken any medicines included in the above definition one-or-more times. The age of onset of psychotherapeutic medicine use was defined as the age at which a respondent first used a psychotherapeutic medicine.

Illicit drug use was defined as any use of drugs to get high, or without a prescription, or in greater quantity than was prescribed. Drug categories included marijuana, amphetamines, barbiturates, tranquilizers, cocaine, heroin, methadone, opiates, and psychedelics, with age of onset of illicit drug use defined as the reported age at which a respondent first engaged in illicit drug use. Age of onset of illicit drug use was distinguished from age of onset of psychotherapeutic medicine use by means of specific decision rules (see Data Analysis).

Data on psychiatric symptoms were obtained using the Diagnostic Interview Schedule. The term symptom was used to denote either the presence of psychiatric symptoms or a discrete spell of psychiatric disturbance occurring at any time in the respondent's life. Symptoms included were nervousness, delusional thoughts, hallucinations, pathologic worry, compulsions, and phobias. Spells included depressive spells, manic spells, and anxiety and panic attacks. The age of onset of a psychiatric symptom or spell was defined as the reported age at which a symptom or spell was first experienced. For individuals experiencing more than one category of symptoms, age of onset was defined as the youngest age at which the respondent reported onset of a symptom.

Data Analysis

Analyses were conducted using Cox proportional hazards models for censored survival data, with illicit drug use and psychiatric symptoms defined as time-dependent covariates.[19–21] Since time (t) in this analysis was defined as chronological age, all models were carefully age-adjusted as suggested by Breslow.[20] Illicit drug use occurring prior to the onset of psychotherapeutic medicine use was coded as zero until the age of onset of such use, at which point it was coded as 1. Thus, for the covariate illicit drug use, a subject would have a value of zero for every age until an illicit drug was used. If an illicit drug was first used at age 18, the subject would have a value of 1 for illicit drug use beginning at age 18 and continuing from that age onward. Psychiatric symptoms were treated in the same way. Subjects who had not used a psychotherapeutic medicine by the time of interview were treated as censored observations.

Forty-five respondents reported the same age of onset for illicit drug use and psychotherapeutic medicine use. This led us to be uncertain about whether prescribed psychotherapeutic medicines had been used illicitly by these respondents. To address this uncertainty we conducted analyses in two ways: first, these individuals were counted as positive for illicit drug use and second, as negative for illicit drug use. There was little difference in the results of these two approaches. In this paper we present the results of the second approach.

In addition to adjustment for age, models were also adjusted for other fixed covariates selected as a result of the heterogeneity with which they were distributed over psychotherapeutic medicine use in the study data, and from evidence from prior research on psychotherapeutic medicine prevalence.[22–28] Fixed covariates were defined as follows, with reference categories in parentheses: race = (Black)/other; sex = (male)/female; marital status = (married)/never married/widowed/separated-divorced; annual household income = <$15,000/$15–19,999/$20–24,999/($25,000+); and help-seeking = yes/(no). Help-seeking was defined as ever having sought help from a professional or a provider for a mental health, nervous or emotional problem, or for problems with alcohol and drugs.

For the purposes of analysis four models were fitted. Model 1 tested for the existence of a predictive association between illicit drug use and psychotherapeutic medicine use in the presence of the fixed covariates. Model 2 tested the hypothesis that the association between illicit drug use and psychotherapeutic medicine use, if present, could be accounted for by the existence of psychiatric symptoms. This was accomplished by entering the term for symptoms into the model, along with illicit drug use and the fixed covariates. Model 3 added an interaction term to Model 2, to test for a possible interaction between psychiatric symptoms and illicit drug use in relation to psychotherapeutic medicine use. Model 4 examined the effect of the sequence of onset of symptoms and illicit drug use on psychotherapeutic medicine use. In particular, Model 4 took the form:

$$h(t) = h_o(t)\exp\{b_1x_1(t) + b_2x_2(t) + \gamma x_1(t)x_2(t) + \\ b_3x_3(t)\{x_1(t)x_2(t)\} + b_4x_4 + \ldots b_px_p\},$$

with $h_o(t)$ representing the underlying hazard of psychotherapeutic medicine use at age t for the reference category (i.e., all covariates = 0). The time-dependent covariates were represented by $x_1(t)$ = illicit drug use, $x_2(t)$ = psychiatric symptoms, $x_1(t)x_2(t)$ = interaction term for (symptoms) (illicit drug use), and $x_3(t)\{x_1(t)x_2(t)\}$ representing those for whom symptom onset preceded illicit drug use, with the term $x_3(t) = 1$ if symptom onset preceded illicit onset, and the fixed covariates were represented by $\{x_4 \ldots x_p\}$. The program BMDP 2L was used for fitting the models.[29]

Models 1–4 were generated for all respondents with complete information on the fixed covariates (n = 1,537). The 25 widowed individuals were excluded from the analyses, since there were too few to generate stable estimates; most exclusions (142 of 179) were due to missing income data.

Since marital status, income, and help-seeking all were measured at the time of interview, the values for these variables prior to the onset of psychotherapeutic medicine use were not known. Therefore, no conclusions about the direction of their association with psychotherapeutic medicine use, if any, can be made. Despite this potential misspecification, these models offer a conservative test of the effect of illicit drug use on psychotherapeutic medicine use, since the covariates in question may be accounting inappropriately for variation due to illicit drug use.

Results

For the entire sample (n = 1,716), 554 (32 percent) of the respondents reported that they had used a psychotherapeutic medicine, 668 (39 percent) were positive for illicit drug use, and 785 (45.7 percent) had experienced psychiatric symptoms. Respondents were defined as positive for the predictor variables if they had used illicit drugs or experienced symptoms prior to the age of onset of psychotherapeutic medicine use or prior to the age at interview.

The age of onset of illicit drug use had a fairly limited range, with onset rarely occurring before age 10 or after age 29. The modal ages for initiation of illicit drug use were 16–19 years of age. Overall, the onset of psychotherapeutic medicine use was concentrated in the older ages of this sample. Instead of a relatively narrow period of initiation as was seen for illicit drug use, the onset of psychotherapeutic medicine use occurred in fairly steady proportions from ages 18–34.[30] The mean age of onset for illicit drug use was 19.35 years, and for psychotherapeutic medicine use 23.88 years.

Table 1 presents the results of the proportional hazards models which included illicit drug use and psychiatric symptoms, with adjustment for age, and for the fixed covariates race, sex, marital status, household income, and help-seeking. Model 1 showed that illicit drug use was associated with an increased risk of psychotherapeutic medicine use (rel. hazard = 2.65; 95% CI = 2.17, 3.21). That is, antecedent illicit drug use was associated with a risk of psychotherapeutic medicine use that was 2.65 times higher than the risk for individuals with no antecedent illicit drug use. In Model 2, the strength of illicit drug use as a predictor of psychotherapeutic medicine use was unchanged by adjustment for psychiatric symptoms. Thus, the presence in the model of a term for symptoms did not alter the relationship between illicit drug use and psychotherapeutic medicine use.

According to Model 3, illicit drug use was associated with an increased risk of psychotherapeutic medicine use whether symptoms had occurred (rel. hazard = 3.60) or had not occurred (rel. hazard = 3.10), with no significant interaction effect. Once again, illicit drug use was a predictor of psychotherapeutic medicine use and symptoms did little to alter this relationship.

Model 4 indicated that when both symptoms and illicit drug use occurred, respondents for whom symptoms preceded illicit drug use were more likely to initiate prescribed psychotherapeutic medicine use than persons for whom illicit drug use preceded symptoms (rel. hazards = 3.86 and 2.64, respectively). However, the confidence intervals were large and overlapped considerably (p = .063). In addition, those individuals with illicit drug use and no symptoms had a risk of psychotherapeutic medicine use similar to those experiencing either sequence of onset of symptoms and illicit drug use, suggesting that the presence of symptoms conferred little, if any, additional risk of psychotherapeutic medicine use over and above illicit drug use.

An additional version of Models 1 and 2 was generated that included onset of alcohol intoxication as a control variable. Onset of alcohol intoxication was defined as the age at which a respondent first reported becoming drunk, and was divided into three dummy variables that were not time-dependent: 1) onset from ages 3–15 years; 2) 16–20 years; and 3) >20 years, with the reference category defined as no onset of alcohol intoxication.

In Model 1, which was adjusted for onset of alcohol intoxication, there was little change in the strength of illicit drug use as a predictor (Table 2), as compared to the unadjusted Model 1 (Table 1). Similar findings were present for the alcohol-adjusted Model 2 (Table 2) as compared to the unadjusted Model 2 (Table 1).

Discussion

All of the models tested in this study showed an

TABLE 1—Cox Proportional Hazards Models of the Association between Psychotherapeutic Medicine Use (1981) and Time-dependent Covariates Illicit Drug Use and Psychiatric Symptoms in Baltimore, with Time Defined as Chronological Age (n = 1,537)

Variables	Model 1 (unadjusted for symptoms)	Model 2 (adjusted for symptoms)	Model 3 (interaction)	Model 4 (sequence of onset)
Illicit Drug Use				
Yes	2.65[*] (2.17, 3.21)	2.61 (2.15, 3.18)		
No	1.00	1.00		
Symptoms				
Yes		1.43 (1.18, 1.74)		
No		1.00		
Illicit Drug Use/Symptoms				
No/No			1.00	1.00
Yes/No			3.10 (2.33, 4.03)	3.07 (2.33, 4.03)
No/Yes			1.60 (1.26, 2.02)	1.60 (1.26, 2.02)
Yes/Yes			3.60 (2.69, 4.80)	
illicit first				2.64 (0.91, 7.72)
symptoms first				3.86 (1.79, 8.34)

[*] Relative hazards are adjusted for age and for fixed covariates sex, race, marital status, household income, and help-seeking. 95% confidence intervals in parentheses.

TABLE 2—Alcohol Intoxication-adjusted Cox Proportional Hazards Models of the Association between Psychotherapeutic Medicine Use (1981) and Time-dependent Covariates Illicit Drug Use and Psychiatric Symptoms in Baltimore (n = 1,528), with Time Defined as Chronological Age*

Variables	Relative hazards Model 1 (unadjusted for symptoms)	Relative hazards Model 2 (adjusted for symptoms)
Illicit Drug Use		
Yes	2.52 (2.04, 3.14)	2.52 (2.02, 3.11)
No	1.00	1.00
Symptoms		
Yes		1.41 (1.15, 1.71)
No		1.00
Alcohol Intoxication Age of Onset		
3–15 Years	1.28 (0.92, 1.79)	1.24 (0.88, 1.72)
16–20 Years	1.08 (0.80, 1.44)	1.05 (0.78, 1.41)
> 20 Years	1.10 (0.79, 1.54)	1.09 (0.78, 1.53)
Never Used	1.00	1.00

*Relative hazards are adjusted for age and for fixed covariates sex, race, marital status, household income, and help-seeking. 95% confidence intervals in parentheses.

association between illicit drug use and the later initiation of prescribed psychotherapeutic medication, even when controlling for initial occurrence of psychiatric symptoms. Symptoms did not change this relation. Model 4 showed that the risk of prescribed psychotherapeutic medicine use was elevated, whether illicit drug use preceded or followed onset of symptoms. Finally, adjustment of the models for the onset of alcohol intoxication did not alter these results.

Learned behavior[15] and sensation-seeking trait explanations[16] are both consistent with the finding that the association between illicit drug use and psychotherapeutic medicine use occurred independent of psychiatric symptoms, as drug use motivated by these explanations would take place without regard to distress state.[31] In terms of maturation and deviance,[13,14] the change in substances used, which may occur from adolescence to young adulthood, could be viewed as a move from behavior that is more deviant to a behavior that is more consistent with adult social roles.[32]

The results of this study are congruent with those of Yamaguchi and Kandel[6] in terms of the association between illicit drug use and psychotherapeutic medicine use, even though our study included a more diverse sample ranging in age from 18–44 years. Whether these results would hold elsewhere or at other times is a question of generalizability best addressed by replication and by greater attention to members of the population not typically interviewed in field surveys. As an example of how results might not generalize to other times, data were gathered in 1981, midway through the cocaine epidemic in the US. The possible impact of the epidemic on our findings is uncertain.

The key limitations to these findings relate to the retrospective character of the study data. For example, validity of self-reports of onset of psychological symptoms may be affected by many factors,[33,34] including severity of the symptoms, length of time since the symptoms were first experienced, and the number of episodes experienced over the lifetime. Similarly, the recency of use of the medicines and the amount of use,[35] as well as the indications for use (i.e., if nonpsychiatric) also could affect the accuracy of the findings.

In addition, the validity of the order in which drugs were used also is relevant. Evidence[36] supporting validity of self-reported order of drugs used comes from a series of repeated measures design surveys of drug use among high school seniors. The findings of these surveys consistently show that marijuana use precedes the use of other illicit substances. These surveys also corroborate findings of two studies[3,4] of sequencing of drug use which used retrospective lifetime reports. Nevertheless, some errors in this regard are possible, as are errors in ordering of onset of symptoms relative to drug use.

In conclusion, this study has shown a predictive association between illicit drug use and psychotherapeutic medicine use, independent of psychiatric symptoms. Future study of these relationships should include: 1) a design which would minimize memory/reporting errors, and 2) a more detailed examination of the etiologic mechanisms that might be responsible for this association.

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Independence Foundation Grant to Establish Nursing Chairs at Nine Universities

Nine of the nation's nursing schools will benefit from a $9 million grant to endow an Independence Chair for Nursing at each institution. The award—made by the Independence Foundation of Philadelphia—is the largest foundation grant in the history of nursing education.

Robert A. Maes, president of the Independence Foundation, said that the interest from the grant should allow nursing programs to engage in long-term planning that will ensure the future of nursing education. "Unlike medical programs," he said, "nursing programs have not been richly endowed and consequently funds for planning have been scarce." The grant from the Independence Foundation will increase the number of nursing chairs in the United States from 38 to 47. Recipients of the grants are nursing schools at the following institutions: Case Western University, Emory University, Johns Hopkins University, New York University, Rush University, University of Pennsylvania, University of Rochester, Vanderbilt University, and Yale University.

The Independence Foundation, established in 1961 in Philadelphia, originated from the Donner Foundation of New York and, over the years, has largely given to secondary education. Since 1985, the Foundation has focused on the national nursing shortage and, two years later, began extensive scholarship funding of each of the nine schools involved in the new grant. In a further effort to draw attention to the national nursing shortage, the Foundation recently announced an increase in those scholarships, as well as the generous endowment of nine chairs for nursing in universities across the nation.

Increasing Rubella Seronegativity Despite a Compulsory School Law

TIMOTHY R. SCHUM, MD, DAVID B. NELSON, MD, MSC, MATTHEW A. DUMA, MD, AND GERALD V. SEDMAK, PHD

Abstract: To determine if lower rubella susceptibility persisted five to seven years after immunization legislation, we retrospectively reviewed the serologic status of 341 outpatients from 1985 to 1987 in an inner-city school age population. Seronegative rates increased significantly during the two-year study period from 4.2 to 24.5 percent (17 percent overall). ($\beta = 6.8\%$, 95% CI = 3.3, 10.3).

Charts were reviewed for 57 of 58 seronegative and 114 seropositive controls. Estimates were then made to the population of 341 subjects. Those with documented rubella immunization had a seronegative rate of 13 percent compared to 19 percent if the immunization status was unknown. For patients who received care in our clinic for < two years, 32 percent were seronegative compared to 10 percent for those treated ≥2 years (odds ratio = 0.31, 95% CI = 0.16, 0.60). Among patients immunized in 1977 or 1978, 33 percent were seronegative compared to 7 percent immunized at other times (OR = 0.21, 95% CI = 0.07, 0.63).

Factors associated with increased rubella seronegativity include immunization in 1977 or 1978 and lack of continuity of care in our clinic. Much of the increase remains unexplained. (*Am J Public Health* 1990; 80:66–69.)

Introduction

In an attempt to prevent congenital rubella syndrome (CRS), rubella vaccination of all 15-month-olds is recommended.[1] This strategy has interrupted the six-nine year cycle of rubella epidemics and cases of rubella have decreased by 98 percent since the prevaccine era of 1969.[2] Outbreaks of rubella still occur in the adolescent/young adult age group, however,[3–5] leaving a serious potential for CRS.

The reported prevalence of rubella susceptible adolescents varies from 7 to 24 percent.[6–9] Nelson, *et al*, showed a dramatic decrease in rubella susceptibility in inner-city adolescent and preadolescent females from 22 percent to 5 percent when a mandatory school immunization law went into effect.[10]

The purpose of the present study was to determine whether this improved serologic status persisted, and what factors affect rubella immune status.

Methods

The Downtown Health Center is an inner-city clinic in Milwaukee, Wisconsin providing general pediatric care. The patient population is 75 percent Black and 85 percent Medicaid. All girls at or near the age of 10 years presenting for well child care, or for school or sports physicals, are routinely screened for rubella susceptibility. We reviewed the results of these tests from August 1985 through July 1987 and compared them with tests done in 1980 and 1981 (the same population with clinic name changed).

On the initial screen, we collected each patient's name, age, date, and result of her rubella titer. All rubella hemagglutination inhibition (HI) titers were measured by the City of Milwaukee virology laboratory using kaolin treatment of serums and one-to-three-day-old chick red blood cells. Titers of <1:10 were considered seronegative.[11]

In an effort to identify factors associated with increased seronegativity, we performed a case control study. From the population of 341 patients with known rubella titers, we did a detailed chart review on a subset of 171 patients—57 of 58 seronegative patients (one chart was unavailable) and 114 seropositive patients. Two seropositive controls were randomly selected for each seronegative case. Information collected from these charts included history and date of rubella immunization, type of immunization, clinic at which immunization was administered, type of health insurance, length of residence in Milwaukee, and current zip code. Seronegative rates for the population of 341 based on time periods could be calculated directly. However, population seronegative rates based on other parameters (history of immunization, duration of care in our institution, year immunized, and clinic of immunization) could not be calculated directly. Estimates of these rates were calculated using the overall incidence rates, odds ratios, and relative frequency of risk factors in the seropositive and seronegative groups.[12] This estimate takes into account the proportion of the subset exposed to a particular variable, e.g., year immunized, and then provides an estimate for the population of 341.

Patients were considered to have a positive history of immunization if a specific immunization date was recorded in their medical record. Immunizations given in our clinic also had the day and site of injection documented.

Differences in rates were measured using the Mantel-Haenszel estimate of the relative risk (odds ratio) with test-based confidence intervals.[13] To control for possible confounding, logistic regression was used.[14] Time trends were determined using a test for linear trend in proportions, a method similar to linear regression except each proportion is weighted by the sample size on which it is based.[15]

Results

Three hundred forty-one patients had their rubella serologic status determined during the two-year study period. Overall, 58 (17 percent) were seronegative. The seronegative rate was 5.8 percent in 1981[10] and 4.2 percent in 1985. However, from 1985 to 1987 there was a consistent increase in rubella susceptible girls to 24.5 percent (Table 1), $\beta = 6.8\%$ (95% CI = 3.3, 10.3).

Subsequent analyses refer to the estimates made on the population of 341 after chart review of the 171 cases. Seventy of 171 cases had documented immunization; 37 of these were immunized at another clinic; 27 were immunized in 1977 or 1978 (Table 2). Children without documented vaccination history were no more likely to be seronegative than children with a documented vaccination history, i.e., 19 percent vs 13

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Editor's Note: See also related letter p 85 this issue.

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TABLE 1—Rubella Immune Status of Females by Six-Month Periods, AUGUST 1985–JULY 1987

	Rubella Status				
	Seronegative*		Seropositive		
Time	Patients	%	Patients	%	Total Patients
AUG 85–JAN 86	3	4.2	69	95.8	72
FEB 86–JUL 86	9	12.7	62	87.3	71
AUG 86–JAN 87	21	21.9	75	78.1	96
FEB 87–JUL 87	25	24.5	77	75.5	102
Totals	58	17.0	283	83.0	341
JUL 80–MAR 81**	15	5.8	242	94.2	257

*β = 6.8% [95% CI = 3.3, 10.3] 1985–87 data.
**Data from Nelson, *et al.*[10]

TABLE 2—Univariate Odds Ratios and 95% Confidence Intervals for Selected Variables and Rubella Antibody Status

		Rubella Antibody Status*		
Variable	N**	% Seronegative	Odds Ratio	95% Confidence Intervals
History of prior immunization				
Yes	70	13	.69	0.36,1.34
Unknown	101	19		
Duration of care in our institution				
≥2 yrs	117	10		
<2 yrs	54	32	0.31	0.16,0.60
Year immunized				
1977–78	27	33	0.21	0.07,0.63
All other years	43	7		
Clinic of immunization				
Other	37	18	0.89	0.31,2.51
DHC***	33	16		

*Estimated

$$\text{Incidence (not exposed)} = \frac{\text{Incidence (total population)}}{[(\text{Proportion exposed}) \times \text{Odds ratio}] + (\text{proportion unexposed})}$$

Incidence (exposed) = Odds Ratio × Incidence (not exposed)

**N = 171 cases with secondary chart review
***Downtown Health Center, Milwaukee, Wisconsin

percent, respectively (Table 2). Patients immunized elsewhere had seronegative rates similar to those immunized at our clinic, 18 percent vs 16 percent, respectively. Children immunized at times other than 1977 or 1978 had lower seronegative rates than children immunized during 1977 and 1978, 7 percent vs 33 percent. Of the 171 patients, 117 were treated at our institution for greater than two years and showed a seronegative rate of 10 percent compared to a 32 percent rate for those treated here less than two years.

Of the 20 seronegative girls with documented immunization histories, 12 were immunized at 15 months or later, four between 12–14 months, and four at or near their first birthday.

Results from a logistic regression analysis are shown in Table 3. Only the following variables were simultaneously entered into the regression model: duration of treatment in our institution, clinic of vaccination, year vaccinated, and history of vaccination. Positive associations with rubella seropositivity were found for treatment in our clinic ≥2 years and for vaccination in years other than 1977–78.

The mean age for serosusceptible girls (11.9) was not significantly different from seroimmune girls (12.4). Type of immunization, e.g., combined measles-mumps-rubella, geographic location, insurance type, and length of time from immunization were not associated with rubella serologic status.

TABLE 3—Multiple Logistic Regression Analysis on Rubella Status

Variables	Odds Ratio	95% Confidence Intervals
Year vaccinated	0.17	0.06, 0.53
Clinic	0.75	0.24, 2.37
Time treated in our institution	0.27	0.15, 0.48
History of prior immunization	0.94	0.20, 7.80

Discussion

The sudden increase between 1985 and 1987 in seronegativity remains largely unexplained. But it is clear that improved rubella seropositivity associated with the school immunization law was not sustained in our population. The 17 percent prevalence of seronegativity in our 1985–87 population is similar to other reports,[7,8] in particular the 22 percent Nelson described prior to the institution of a statewide school immunization law.[10] Since July of 1987, the seronegative rate in our clinic has leveled off at 18 percent (54 of 300 from August 1, 1987 through January 31, 1989). The Milwaukee Health Department also noted an increase in rubella susceptibility in 1987, the highest level since 1979 (Table 4).

Several factors might explain increased seronegativity. One possible explanation would be failure to enforce the school immunization law. The school system, however, has reported no change in enforcement practices. Documentation of actual dates of immunization are and have been required for school entry. Another possible explanation is patient mobility and lack of continuity of care. The increased seronegativity in patients who have been attending our clinic for less than two years suggests this might have occurred. However, strict enforcement of the law should have identified these children.

Perhaps our 17 percent seroprevalence is biased or not generalizable because we could document immunization history in only 70 of 171 charts reviewed (41 percent). Seronegativity was 5.7 percent in Nelson's study even though they could document immunization history in only 46 percent of their patients,[10] making this an unlikely explanation.

Primary or secondary vaccine failure could explain the increased seronegativity, but our rate of 17 percent is higher than reported rates of 0–12 percent.[16–18,24,25] Secondary vaccine failure is unlikely as we found no association between seronegativity and time elapsed since immunization.

There was no evidence that the HI testing methodology explained the increased seronegativity. The HI test method, laboratory personnel, and titer cutoffs (1:8–1:10) have not changed. Newer testing methods are available and are more sensitive at detecting low antibody titers[21–23] but the standard for comparison still remains the HI test.

The current vaccine was licensed in 1979. We found significantly increased seronegativity for those immunized in 1977–78 but not before or after that time period. Perhaps there was a problem with particular lots of vaccine in those two years. Although problems with improper handling have been reported,[24,25] this is an unlikely explanation because seroprevalence rate was the same regardless of the clinic site.

Finally, the differences found might relate to selectively screening girls who are likely to be seronegative. This, however, is unlikely. The clinic clientele has not changed over the last 10 years. We see mostly Medicaid HMO (health maintenance organization) patients living in the inner-city and our referral patterns have not changed. All females age 8 or over presenting for well child care are eligible for screening. The vast majority are screened, although a few, known to be unvaccinated, will be immunized and not screened. Most patients are unaware of their rubella status and we do not require immunization records prior to screening. There is no reason to suspect this population was at increased risk for rubella seronegativity as most came to the clinic just for routine care rather than being referred for a specific problem.

TABLE 4—Rubella Serology from Specimens Submitted to the Milwaukee Health Department 1979–87

	Rubella Status		
	Seronegative	Seropositive	Total
Year	No. (%)	No. (%)	No.
1979	666 (13.7)	4190 (86.3)	4856
1980	472 (13.4)	3043 (86.6)	3515
1981	304 (12.8)	2071 (87.2)	2375
1982	114 (10.4)	985 (89.6)	1099
1983	140 (10.5)	1194 (89.5)	1334
1984	117 (11.6)	890 (88.4)	1007
1985	123 (12.9)	832 (87.1)	955
1986	107 (12.2)	768 (87.8)	875
1987	189 (22.0)	671 (78.0)	860
Totals	2232 (13.2)	14644 (86.8)	16876

Despite the increased seronegativity, we have not had sustained rubella transmission in Milwaukee since 1977 when a university-associated outbreak occurred. However, there was a mumps outbreak in the general population in 1987[26] and the possibility of a rubella outbreak exists. If the rate of serosusceptibility to rubella remains as it is, then a large number of women from this population may be entering their childbearing years, susceptible to rubella.

In summary, the dramatic impact of a mandatory school immunization law was not sustained seven years later. Rubella seroprevalence is now the same as it was prior to mandatory immunization laws although reported school enforcement has remained the same. Patients without a documented history of rubella immunization should be immunized as per current recommendations.[1] Females with prior rubella immunization should have rubella serology performed to assure immunity as documented history of rubella immunization did not guarantee seropositivity in our study.* The lack of continuity of care and rubella immunization during 1977–78 were significantly associated with increased seronegativity in our population. The sudden increase in seronegativity in just two years from 4 to 24 percent still remains largely unexplained but has leveled off at 17–18 percent. School enforcement of mandatory immunization laws is important but in Milwaukee is not sufficient to assure that girls approaching adolescence are protected against rubella.

ACKNOWLEDGMENTS

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*See Letters section for recommendations of the Centers for Disease Control on this point.

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Noise and Hearing Loss Is Focus of Upcoming NIH Consensus Development Conference

Noise and hearing loss is the subject of an upcoming consensus development conference, to be held January 22-24, 1990 at the Masur Auditorium of the National Institutes of Health Clinical Center, Bethesda, Maryland. In the background material provided by NIH, it was noted that:

- Approximately 28 million individuals in the United States have impaired hearing;
- Exposure to occupational and recreational noise has, in many instances, caused irreversible damage to the inner ear;
- Sources of potentially damaging noise include high-volume music, recreational vehicles, airports, heavy traffic, lawn mowers, household appliances, chain saws, and power tools, to name a few;
- Hazardous levels of noise on the job are experienced daily by firefighters, military personnel, disc jockeys, construction workers, musicians, farmers, factory workers, taxicab/bus/truck drivers, and others.

The NIH conference will bring together specialists in hearing, physiology, and pathophysiology and other relevant fields. The purpose of the conference is to reach agreement on noise and hearing loss. Experts will present current scientific thinking about the diagnosis, management, and prevention of hearing loss, and concerned voluntary organizations will be invited to make statements. On the basis of the scientific evidence presented, the consensus panel will prepare a draft report and invite comments.

Key questions to be addressed include: 1) What is noise-related hearing loss? 2) What sounds can damage hearing? 3) What factors, including age, determine an individual's susceptibility to noise-related hearing loss? 4) What can be done to prevent noise-related hearing loss? 5) What are the directions for future research.

To register for the noise and hearing consensus conference or to obtain further details, contact: Dina Rice, Prospect Associates, Suite 500, 1801 Rockville Pike, Rockville, MD 20852. Telephone: (301) 468-6555.

Public Health Briefs

Foreign-Born and US-Born Black Women: Differences in Health Behaviors and Birth Outcomes

HOWARD CABRAL, MPH, LISE E. FRIED, MSPH, SUZETTE LEVENSON, MPH, HORTENSIA AMARO, PHD,
AND BARRY ZUCKERMAN, MD

Abstract: We studied health behaviors and birth outcome among 201 foreign-born and 616 US-born Black women receiving prenatal care at Boston City Hospital. Foreign-born women had better pre-pregnancy nutritional status and prenatal health behaviors, and their infants had greater intrauterine growth. Black women are not a homogeneous group; culture and ethnicity, in addition to other variables, must be considered in the study of their birth outcomes. (*Am J Public Health* 1990; 80:70–72.)

Introduction

In the United States, Black neonates are at high risk of low birthweight and increased perinatal mortality.[1–3] The rate of low birthweight among foreign-born Blacks is, however, two-thirds that of US-born Blacks.[4] The percentage of Black births in the US to mothers of foreign birth has increased in recent years. By 1986, this percentage had risen to 8 percent, an increase of 64 percent from 1976 and 39 percent from 1980, largely in urban areas of the northeast US.[5–7]

Differences in the rates of low birthweight among Blacks by foreign birth status may be related to lifestyle, since the health behaviors of those born outside of the US are different from those born in the US. For example, among Hispanics, those born outside of the US are less likely to smoke cigarettes, use illicit drugs, or receive late or no prenatal care than Hispanics born in the US, whose health behaviors more closely resemble those of US natives of other ethnic groups.[8–12,*]

The goal of this paper is to ascertain whether similar differences in health behaviors, and consequently birth outcomes, are present in a sample of 201 foreign-born and 816 US-born Black women receiving prenatal care at Boston City Hospital.

Methods

As part of a study of maternal health, psychosocial characteristics, health behaviors, and pregnancy outcome initiated in 1984, subjects were consecutively recruited in the prenatal clinics of Boston City Hospital. Interviews were conducted in English or Spanish at the time of prenatal registration and in the immediate postpartum period. Information on each participant's race and country of birth was obtained from closed-ended items on the initial interview. Substance use was assessed by self-report, except for marijuana and cocaine, which were also identified through urine assays. Infant assessments were conducted by pediatricians blind to the hypothesis of the analysis.

The 817 Black women presented here are a subsample of the total study sample of 1,226 women who delivered a liveborn, singelton infant and on whom complete data, including two interviews, a maternal medical record review, and an infant assessment, were available. A more detailed description of the study protocol and population has previously been published.[13]

Results

The demographic characteristics of the 817 Black women presented here are comparable to those of the low-income (47 percent with a monthly income of less than $500, 51 percent currently receiving Medicaid), inner-city prenatal population served by Boston City Hospital. Eighty-two percent of these women were unmarried, 15 percent were 18 years of age or younger, and 36 percent had received less than 12 years of education. Of the 616 US-born women, 7 percent were first generation Americans. Of the 201 foreign-born Black women, 72 percent were from Caribbean countries (40 percent from Haiti, 11 percent from Jamaica, 21 percent from other Caribbean countries), 17 percent from Africa, 8 percent from Cape Verde, and 3 percent from England. Country of origin was not significantly associated statistically with the dependent variables used in this analysis among women of foreign birth.

Foreign-born women were more likely to be older, married, better educated, to have better pre-pregnancy weight-for-height, and to have had nine or more visits for prenatal care than those born in the US (Table 1). The two groups of women were similar on parity, trimester of first prenatal visit, and pregnancy weight gain of less than 16 lbs. They were also less likely to smoke cigarettes, drink alcohol, or use marijuana, cocaine, or opiates during pregnancy (Table 2); in addition, they used fewer cigarettes and consumed less alcohol (data not shown).

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[*] Amaro H, Whitaker R, Coffman G, Heeren T: Acculturation and Drug Use: Findings from the Hispanic HANES. Paper presented at the First National Conference of the National Hispanic Psychological Association, New York, August 1987.

TABLE 1—Comparison of Foreign-Born US-Born Black Women on Demographic and Selected Obstetric Factors

Factors	Foreign-Born (n = 201)	US-Born (n = 616)	% Difference	95% CI of Difference
	%	%		
Maternal Age				
18 years or younger	7	18	−11	(−6, −16)
Marital Status				
Married	33	12	21	(14, 28)
Maternal Education				
11th grade or less	22	40	−18	(−11, −25)
Monthly Household Income*				
$500 or less	44	48	−4	(−13, 5)
Pre-Pregnancy Weight-for-Height				
Less Than 90% of Ideal	6	13	−7	(−3, −11)
Primiparity	47	47	0	(−8, 8)
Trimester of First Prenatal Visit				
First or Second	87	85	2	(−3, 7)
Prenatal Care				
Less than nine visits	33	47	−14	(−6, −22)
Pregnancy Weight Gain				
Less than 16 lbs	12	16	−4	(−9, 1)

* of the 81% reporting an income

TABLE 2—Comparison of Foreign-Born US-Born Black Women on Substance Use Anytime During Pregnancy

Substance Used	Foreign-Born (n = 201)	US-Born (n = 616)	% Difference	95% CI of Difference
Cigarettes	12%	53%	−41	(−35, −47)
Alcohol	45%	64%	−19	(−11, −27)
Marijuana*	6%	39%	−33	(−28, −38)
Cocaine*	3%	24%	−21	(−17, −25)
Opiates**	1%	4%	−3	(−1, −5)
Other Drugs	1%	2%	−1	(−3, 1)

*Determined by positive self-report and/or urine assay.
**Heroin, Methadone

As seen in Table 3 infants of the foreign-born women were, at birth, heavier, longer, had larger head circumferences, and had longer durations of gestation than did infants of the US-born women. Foreign-born Black women in this sample were also less likely to have given birth to a low birthweight (≤2500 grams) infant (odds ratio = 0.59, 95% CI = 0.33, 1.03) or a premature (<37 weeks gestation) infant (odds ratio = 0.46, 95% CI = 0.22, 0.94) than those born in the US.

After controlling for the effects of gestational age, weight gain during pregnancy, pre-pregnancy weight-for-height, marital status, maternal age, level of education, the number of visits made for prenatal care, and the use of cigarettes, alcohol, marijuana and cocaine during pregnancy, infants born to foreign-born Black women remained heavier, longer, and larger in head circumference at birth than infants born to US-born Black women (Table 3). Duration of gestation was not different between infants of foreign-born Black women and infants of mothers born in the US. The adjusted odds of having a low birthweight infant or a premature infant remained decreased for foreign-born Black women as opposed to Black women born in the US (odds ratio, low birthweight = 0.81, 95% CI = 0.42, 1.53; odds ratio, prematurity = 0.54, 95% CI = 0.24, 1.18).

TABLE 3—Differences in Fetal Growth Measured at Birth between Infants of Foreign (N = 201) and US-Born (N = 616) Black Women

	Foreign-Born vs US-Born		Foreign-Born vs US-Born	
	Unadjusted Difference	95% CI of Difference	Adjusted Difference	95% CI of Difference
Birthweight	272 g	(174, 369)	135 g	(51, 218)
Length	1.3 cm	(0.9, 1.8)	0.7 cm	(0.2, 1.1)
Head Circumference	0.8 cm	(0.5, 1.2)	0.4 cm	(0.1, 0.7)
Gestational Age	0.4 wk	(0.01, 0.7)	0.3 wk	(−0.1, 0.7)

Discussion

The infants of foreign-born Black women experienced greater intrauterine growth than those of Black women born in the US, even after accounting for many of the factors known or suspected to influence fetal growth. Why such a difference remained deserves more in-depth study than these data can provide.

The following limitations should also be considered when interpreting these findings:

- First, although the percentage of births to foreign-born Black mothers in these data virtually matches that of the entire northeast US (26 percent) in 1986,[7] it is possible that the characteristics of the women studied here are unique to the Black population who use Boston City Hospital as their primary source of prenatal care.
- Second, foreign birth status is an approximation of the cultural characteristics of these women and does not necessarily reflect the heterogeneity of either the US or foreign-born Black women in this regard.
- Third, the exclusion of non-English or non-Spanish-speaking women, who may have been less acculturated to US norms than those who participated, may have affected these findings.

These data illustrate some important differences in prenatal risk and birth outcome by foreign birth status within the prenatal population of Black women at Boston City Hospital. They suggest that low-income Black women are not a homogeneous group in terms of prenatal risk and birth outcome, and highlight the importance of considering ethnicity as indicated by place of birth, in addition to race, as a demographic factor in the study of the relationship of maternal prenatal health behaviors and birth outcomes.

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Some Sources of Error in the Coding of Birth Weight

ANDREW J. BRUNSKILL

Abstract: Three types of error in coding birth weight to computer tapes are described: 1) confusion of ounces with pounds, 2) mistaken reading of one pound as eleven pounds, and 3) errors in placement of the decimal. All will allocate low-birth weight infants to high birth weight categories. Examination of the reported gestational age of the infant or of the reported cause of death may allow these errors to be detected. (*Am J Public Health* 1990; 80:72–73.)

Introduction

There is increasing use of computerized birth certificate records in pediatric epidemiology. Such records may contain both systematic and random errors. David described sources of systematic error in the recording of birth weight and gestational age in North Carolina birth certificates.[1] In that study, it was suggested that when birth weight was abnormally large for gestational age, gestational age was most likely to be inaccurate. However, original birth certificates were not actually compared to the birth tapes for accuracy of transcription. Although the need for careful editing of data is emphasized by experienced researchers, the nature and the effects of errors which occur in coding seem to have received less emphasis.

Methods

This analysis is based on a partial review of the cases present in a state birth tape of 1984–85 data. For those infants who had died within the first year of life, linked death certificates were also available. All White infants with birth weights of 4.5 kgs or greater ("macrosomic") were sampled and a corresponding random sample in an approximately 2:1 ratio of infants with birth weights 2.5–4.0 kgs. This selection, which related to the primary aims of a project on macrosomia, yielded 5,399 macrosomic and 10,250 control infants. Within these samples, the records of all those infants who had died in the first year of life were reviewed and the birth weight reported on the birth certificate compared with the birth weight reported on the computer tape. Additionally, 34 birth certificates of those infants with the highest birth weights who had not died were reviewed to bring the total number of birth certificates reviewed to 120.

Results

Within the high birth weight group, 39 infants were reported to have died within the first year of life. However, 18 cases of miscoded birth weight were detected in this group, leaving only 21 infants whose birth weights were confirmed as being high. Within the normal control group, 47 infants were reported to have died within the first year of life. The birth weights of all these infants were confirmed as correctly coded. All of the additional high birth weight infants who had survived the first year of life and whose certificates were checked had correctly coded birth certificates.

Three types of errors appeared to be responsible for the miscoding:

- One type of error is the confusion of ounces with pounds. Examples include 12 oz being coded as 5,443 grams (12 lbs), and 14 oz being coded as 14 lbs. This error occurred in nine of the 18 cases of detected coding error.
- The second type of error consisted of mistaken reading of birth weights of one pound or more as eleven pounds. The ease with which one pound can be misread as eleven pounds can be understood by considering the abbreviation used for one pound (1 lb) where the lower case letter "l" is mistaken for a second digit one "1." This is illustrated in Figure 1 which consists of a photocopy of the relevant part of the six birth certificates in which all six birth certificate weights were coded as eleven pounds. In fact, only the infant whose birth certificate is coded as "e" had a true weight of eleven pounds. This error occurred six times in the series.
- The third set of errors occurs when a decimal is wrongly placed in metric values. This includes cases such as 510 grams being mistakenly coded as 5100

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Infant Formula Promotion and the Health Sector in the Philippines

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Abstract: We use data collected in 1984, 1986, and 1988 from public and private health facilities in the Metropolitan Cebu area of the Philippines to assess effects of World Health Organization and Government of the Philippines Codes of Marketing of Breast Milk Substitutes on distribution of formula supplies. Distribution of free and low-cost infant formula declined drastically between 1986 and 1988. Industry compliance was almost complete. (*Am J Public Health* 1990; 80:74–75.)

Introduction

After nearly a decade of controversy on the role of infant formula promotion in changes in Third World breast-feeding behavior, the World Health Assembly adopted the International Code of Marketing of Breast Milk Substitutes on May 21, 1981. Most troublesome among these provisions was Article 6, paragraph 6, stating that "[d]onations or low-price sales to institutions or organizations of supplies of infant formula or other products within the scope of this Code, whether for use in the institutions or for distribution outside them, may be made. Such supplies should only be used for, or distributed to, infants who have to be fed on breast milk substitutes."[1]

The World Health Organization (WHO) Guidelines issued on April 10, 1986, defined "infants who have to be fed on breast milk substitutes" as those who have rare metabolic disorders, cannot suck, or are motherless through death or abandonment as well as those whose mothers have decided not to breast-feed, fully or partially, for whatever reason.[2] The World Health Assembly passed a resolution on May 16, 1986 urging governments to take action "to ensure that the small amounts of breast-milk substitutes needed for the minority of infants who require them in maternity wards and hospitals are made available through the normal procurement channels and not through free or subsidized supplies."[3]

On December 22, 1986, the Philippine Government passed its own code of infant food marketing (the GOP code),[4] banning the distribution of free and subsidized formula supplies to hospitals.

WHO in 1988 affirmed that the International Code, which permits free or subsidized supplies of infant formula for infants who have to be fed on breast milk substitutes, had not been modified by the 1986 World Health Assembly Resolution.[5,6]

While the International Code acknowledges a legitimate market for infant formula that includes mothers who cannot breast-feed or choose to bottle-feed, the National Code for Marketing of Breast Milk Substitutes of the Government of the Philippines is silent on the issue of procurement and forbids infant food manufacturers and distributors from giving free supplies to the general public, hospitals, and other health facilities and medical personnel except on request by, or with the approval of, the Department of Health (DOH). Approval by the DOH is given only for catastrophic occurrences such as floods or famine. Violators of the GOP Code face fines or imprisonment. Omission of any reference to procurement in the GOP code raises questions on how to obtain formula for infants unable to breast-feed and/or whose mothers are unwilling or unable to suckle them. We analyzed existing data to assess the effect of these different codes on hospital practices.

Methods

Between 1984 and 1988, as part of a larger study, we surveyed health facilities serving the Philippines' second largest metropolitan area (Cebu City) three times. During this period, the major infant food companies serving this market (American Home Products, Bristol-Myers, Nestlé, and Abbott) had accepted the WHO Code.

We collected data from 73 (1984), 78 (1986), and 72 (1988) Metropolitan Cebu health facilities, all providing prenatal and obstetrical care. These facilities include all public and private hospitals and a representative sample of clinics and health centers. Approximately 70 percent of these facilities are public (government) institutions; about one-fifth of the public facilities and three-fourths of the private facilities have on-site obstetrical services; the remainder provide home-delivery assistance.

We obtained information on whether facilities had received any infant formula, other milk, or food supplies in the three months preceding each survey from manufacturers of those products; the brand name and source of each sample received; how the facilities had used infant formula supplies; and their policies toward industry representatives. The 1988 survey asked additional questions on procurement practices particularly for higher risk infants such as premature babies confined in special care nurseries.

Results

The proportions of facilities receiving as manufacturers' gifts any kind of milk remained basically unchanged between 1984 and 1986 but fell in the following two years (Table 1): infant formula distribution declined by 95 percent, while the proportion of facilities with formula for older infants increased by 80 percent in the same period. These foods intended for use as the liquid part of the weaning diet are not covered by the WHO or GOP codes.* The same applies to any weaning food.

Before passage of the GOP code, public facilities with on-site delivery were most likely to receive free formula (Table 2). All on-site delivery facilities show an increase in

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*Editor's Note: "Follow-up" formulas have not been marketed successfully in the United States. A policy statement of the American Academy of Pediatrics, "Follow-up or Weaning Formulas," appears in the February 1989 issue of AAP News.

TABLE 1—Percentage of Facilities Receiving Any Infant Food as Gifts by Type

Type of Food	1984	1986	1988
Infant formula	57.5	52.6	2.8
Follow-up formula	11.0	15.4	27.8

distribution of formula supplies between 1984 and 1986 by all manufacturers. In the 1988 survey we find only one manufacturer still distributing free formula supplies. Recipients were two private health facilities with on-site delivery which had received a soy-based preparation.

In 1984 and 1986, over half the total sample of facilities had distributed free samples to mothers; in 1988, only 3 percent still did so (data not shown).

The proportions of facilities allowing representatives to visit their patients rose from 30 percent to 40 percent from 1984 to 1986. In 1988, only 2.6 percent of the facilities allowed them to do so. Moreover, only seven of 72 facilities in Metropolitan Cebu still supplied clients' names to industry representatives—a marked decrease from previous years.

The 1988 survey found that seven private facilities sold infant formula in their own hospital pharmacies, an increase from two in 1984 and five in 1986. Bottle-feeding mothers who deliver in these private facilities have a choice of purchasing formula from these pharmacies or from outside sources. Mothers delivering in public hospitals have to buy formula from outside retail outlets if they need or want to bottle-feed their newborns.

Discussion

The trends shown over these three surveys suggest that the WHO International Code of 1981 had little direct influence in reducing the level of infant formula donations to hospitals which the Code did not prohibit.

TABLE 2—Percentage of Facilities Receiving Free Infant Formula Supplies from Specific Industry Sources by Facility Type and Year of Survey[a]

	1984		1986		1988	
Facility Type	%	(N)	%	(N)	%	(N)
Public facility						
With on-site delivery	85.7	(7)	100.0	(7)	0.0	(6)
Without on-site delivery	54.4	(46)	38.0	(50)	0.0	(46)
Private facility						
With on-site delivery	78.6	(14)	93.3	(15)	13.3[b]	(14)
Without on-site delivery	0.0	(6)	16.7	(6)	0.0	(6)

[a] Formula supplies were received within a three-month period preceding each survey.
[b] Soy-based formula only.

While committing itself to eliminating indiscriminate distribution of infant formula supplies, industry took the position that the responsibility of providing directives to health care facilities and manufacturers on infant formula promotion rests with national governments. We observed significant decreases in such activities in 1988 which may be attributable to restrictions and penalties of the National Code passed by the Philippine Government late in 1986.

Whether the changes in the distribution of free supplies observed in the 1988 survey will lead to an increase in breast-feeding among Filipino mothers and an improvement in infant health and development is still unknown. What is apparent is that present policies and practices are designed to make it inconvenient and difficult for mothers to bottle-feed in hospitals.

Finally, it is important to note that this study does not imply that the WHO Code has been of no value. The WHO Code has served as the basis for the development of country codes and has centered extensive publicity on the need to use legislation as one means of promoting breast-feeding. The GOP Code is noteworthy in being one of the first to implement the purpose of the 1986 WHO resolution.

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A Randomized Evaluation of Smoking Cessation Interventions for Pregnant Women at a WIC Clinic

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Abstract: Pregnant smokers attending a local health department WIC clinic were randomly assigned to one of two self-help smoking cessation programs or usual care. The multiple component program resulted in larger quit rates than usual care during the last month of pregnancy (11 percent vs 3 percent) and postpartum (7 percent vs 0 percent). Achieving quit rates in WIC similar to those in studies conducted at prenatal care settings, suggests that smoking cessation programs for low-income pregnant WIC clients are feasible. (*Am J Public Health* 1990; 80:76–78.)

Introduction

Estimates suggest that 32 percent of all women of child-bearing age smoke with only 21 percent quitting during pregnancy.[1] Evaluations of self-help cessation programs have demonstrated some success with pregnant women receiving prenatal care at an HMO (health maintenance organization),[2] private practices,[3] hospital clinics,[4]* and local health departments.[5]

WIC (the Special Supplemental Food Program for Women, Infants, and Children) is an unexplored setting for smoking cessation research. Analyses of matched WIC-birth certificate data for Missouri in 1982 indicated that 45 percent of pregnant women WIC participants were smokers,** 50 percent higher than the 30 percent rate observed among all Missouri women that year. Data on 127,512 pregnant women enrolled in several state WIC programs from 1979 through 1985 suggested that smokers were 2.5 times more likely to deliver a low birthweight infant, and that very few participants stopped smoking following their enrollment in WIC.***

The present investigation is a randomized trial of a self-help smoking cessation program implemented at a local health department WIC clinic.

Method

The study was conducted from August 1985 to August 1986 at the Kent County Health Department in Grand Rapids, Michigan. A brief structured interview was used to characterize an incoming pregnant women as a current smoker (at least one cigarette per day), a nonsmoker, or a former smoker. Current smokers were asked to participate in the study and to complete an informed consent statement and a pretest. Following informed consent, participants were randomly assigned to one of three groups.

The usual care (UC) group received printed information about the risks of smoking during pregnancy and completed the clinic in the traditional manner.

The multiple component (MC) group received a 20-minute one-to-one counseling session which included both risk information and behavior change components. The risk information component employed the *Because I Love My Baby* materials developed by the American Lung Association. These materials included a "flip chart" used by the health educator in presenting the information and a printed brochure given to clients to take home.

The behavior change component of the MC intervention employed a self-help manual adapted from Windsor, *et al*,[5] and from the American Lung Association's *Freedom From Smoking* program. Behavioral contracting and self-monitoring were the primary strategies. An individual behavioral contract was developed during the session which specified a quit date and selection of some significant other as a co-signer. Self-monitoring included charts for recording daily smoking behavior, and the development of an individualized plan of action for breaking recorded behavioral chains.

The risk information (RI) intervention was a face-to-face session of about 10 minutes duration. The health educator used the same "flip chart" as the MC group and provided the factual brochures, but did not present behavior change counseling, or furnish the self-help manual.

During the postpartum WIC visit, the posttest was completed, which included questions about smoking during the final month of pregnancy as well as postpartum. The postpartum self-reports, on average, occurred 4.7 weeks following birth of the infant, a relatively short period. Saliva samples ($N = 66$) were collected from the last one-third of the participants. The correlation between the saliva thiocyanate and self-report measures ($r = .25$) was similar to correlations found in more extensive analyses with larger samples that showed excellent sensitivity and specificity.[6,7] Saliva samples were equally distributed across groups (28 percent of MC participants, 30 percent of RI participants, and 32 percent of UC participants). Resource constraints did not allow collection of more complete data.

Demographic and health service data were abstracted from WIC program records for study participants and for the women who refused to participate. Agreement rates between different coders, and between the WIC record and prenatal care medical record, for a 30 percent random sample, were 99.2 percent and 87.9 percent, respectively.

Results

Two hundred and seventy-one of 641 pregnant women entering the clinic were identified as current smokers (42 percent); 219 volunteered to participate in the study (81 percent). Of these 219, posttest data were collected from 186 (15 percent attrition rate) (about one-fourth actually refused to complete the posttest; three-fourths either moved from the community, left the

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*Secker-Walker RH, Flynn BS, Solomon LJ, Collins-Burris L, LePage S, Mead PB: Attitudes, beliefs and other smokers: Factors affecting smoking cessation in pregnancy. Paper presented at 114th annual meeting of the American Public Health Association, Las Vegas, Nevada, 1986.

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WIC program, or miscarried). Those lost to follow-up were treated as continuing smokers in the following analyses.

Differences between study participants and refusals on variables available from the WIC record were relatively minor for important variables as were study group differences (Table 1).

Table 2 presents the quit proportions and differences by treatment group for both follow-up periods. The MC group quit rate was higher than the UC group at both post time periods. While all UC participants had returned to smoking postpartum, some proportion of the participants in the MC and RI groups remained smoke free.

Discussion

The 42 percent smoking rate among pregnant women entering the WIC clinic was similar to the estimates of Land and Stockbauer,** and was indicative of the important need for smoking intervention at WIC clinics. The quit rate for the MC group (11 percent) was just under that found at prenatal care settings serving low-income women. Both Windsor, *et al*,[5] and Secker-Walker *et al*,* in studies conducted at low-income public maternity clinics, reported quit rates of 14 percent.

Although limitations of the present formative evaluation study include small sample size, and lack of complete biochemical verification, the findings have preliminary implications for extending health education services to pregnant smokers visiting WIC clinics. The high level of program acceptance among our WIC participants (81 percent) is impressive. Although prenatal care settings see pregnant women more regularly, offering the possibility of multiple health education contacts (and potentially improved quit rates), the WIC program, with broader eligibility than typically found at public prenatal care clinics, would seem to offer greater opportunity to reach a larger share of low-income women. Additionally, WIC clinics offer the advantage of a centralized site where health educators and public health nurses can be made available to provide one-to-one counselings.

TABLE 1—Comparison of Multiple Component, Risk Information, Usual Care, and Refusal Groups on Selected Demographic and Health Care Variables

Variables	Multiple Component (n = 72)	Risk Information (n = 70)	Usual Care (n = 77)	Total (n = 219)	Refusals (n = 52)
Mean Age	22.3	23.4	22.4	22.7	23.8
Parity					
% Primiparous	33.8	27.9	23.4	28.2	21.2
% Multiparous	66.2	72.1	76.6	71.8	78.8
Race					
% Black	18.3	25.0	19.5	20.8	18.9
% White	78.9	75.0	71.4	75.0	75.5
% Other	2.8	0.0	9.1	4.2	5.7
Trimester Prenatal Care Began					
% First	88.7	80.3	82.7	84.0	NA
% Second	11.3	18.2	16.0	15.1	NA
% Third	0.0	1.5	1.3	0.9	NA
Trimester of First WIC Clinic Visit					
% First	22.5	27.5	19.5	23.0	23.1
% Second	54.9	44.9	48.1	49.3	50.0
% Third	22.5	27.5	32.5	27.6	26.9
Source of Payment					
% Insurance or HMO	14.5	9.1	21.0	15.1	8.5
% Medicaid	79.0	80.0	71.0	76.5	80.9
% Self-pay or other	6.5	10.9	8.1	8.4	10.6
Received Advice from Physician to Quit					
% Yes	68.6	67.1	68.9	68.2	NA
% No	31.4	32.9	31.1	31.8	NA
Mean Number of Years Smoking	7.2	7.5	7.3	7.4	NA
Mean Age Smoking Began	15.1	15.8	15.1	15.3	NA
Number of Quitting Attempts Prior to Pregnancy					
% None	26.4	34.3	28.9	29.8	NA
% One or more	73.6	65.7	71.1	70.2	NA
Number of Quitting Attempts Following Pregnancy					
% None	69.4	64.3	66.2	66.7	NA
% One or more	30.6	35.7	33.8	33.3	NA
Mean Number of Cigarettes per Day Prior to Pregnancy	19.9	19.7	20.3	19.9	NA

NOTE: All differences between MC, RI and UC groups are not statistically significant at the .05 level using chi-square tests for categorical variables and ANOVA for differences between means. Differences between the total intervention group and the refusal group are also not significant. Sample size may vary slightly due to missing data. NA = not available from WIC record.

TABLE 2—Quit Proportions and Differences at the Ninth Month of Pregnancy and Postpartum

	At Ninth Month	Postpartum
Multiple Component (N = 72)	.111 (.039, .183)	.069 (.010, .128)
Risk Information (N = 70)	.071 (.011, .131)	.071 (.011, .131)
Usual Care (N = 77)	.026 (−.009, .062)	.000 (0.00, 0.00)
Total (N = 219)	.068 (.035, .101)	.046 (.018, .074)

Note: 95% CI in parentheses

At Ninth Month	Difference	95% CI
MC vs UC	.085	(.004, .166)
RI vs UC	.045	(−.025, .115)
MC vs RI	.040	(−.054, .134)
Postpartum		
MC vs UC	.069	(.011, .128)
RI vs UC	.071	(.012, .130)
MC vs RI	.002	(−.082, .086)

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Eight-Year Follow-Up Results of an Adolescent Smoking Prevention Program: The North Karelia Youth Project

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Abstract: In the North Karelia Youth Program five to 10 classroom sessions over two years taught skills to resist pressures to start smoking to 13 to 15 year old students. Compared to students from comparison schools, the treatment groups reported less smoking immediately after the intervention and in a four-year follow-up survey. At the eight-year follow-up, there was consistent evidence of possible preventive effects only among those who had been nonsmokers when the program began. (*Am J Public Health* 1990; 80:78–79.)

Introduction

From 1978 to 1980, a study to demonstrate methods for reducing cardiovascular risk factors in 13-year-old students was carried out in the province of North Karelia, Finland. One of the main aims was to prevent the onset of smoking.[1] Although many studies of the effectiveness of persuasive and behavioral approaches to smoking prevention have been carried out,[2] long-term follow-up is rare.[3,4] Details of the methods and the outcome of the intervention immediately after the program[1], after half a year[5], and two years after the program[6] are reported elsewhere. This brief report describes the follow-up results eight years after the program ended.

Methods

In brief, three pairs of matched schools were chosen for the program. In two "direct" program schools, 10 sessions were carried out during a two-year period from 1978 to 1980 by the project health educator and trained peer leaders. In two "county-wide" program schools, five behavioral sessions were to be carried out by health education teachers at the schools who were trained as part of an effort to disseminate new curricula throughout North Karelia. Two schools were chosen for controls from another province in eastern Finland. In each matched pair, one school was chosen from the capital of the county and the other from a rural village. The program began in 1978 among the seventh graders and finished with the ninth grade in autumn 1980. During the persuasive-behavioral sessions, students were taught about

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social pressures that may lead to smoking (peer, adults, and mass media) and were trained to deal with such pressures through role playing and other active learning techniques. Older peer leaders assisted in implementing the program at school. The short-term and long-term health hazards were also discussed during the program. At the pretest, there were 897 students in the seventh grade in the three pairs of matched schools. At the posttest, the participation rate varied from 92–97 percent, in the eight-year follow-up from 72–79 percent. In all the surveys, the smoking habits of students were asked by the following question: Do you smoke now? Multiple choice responses: 1) Not at all; 2) Less than once a month; 3) One to two times a month; 4) One–two times a week; 5) Daily.

Results

When the youth were in the seventh grade, the proportion who reported smoking at least once a month varied in schools from 4–9 percent. Immediately after the program, there were one-third fewer smokers in both types of intervention schools compared to control schools. One half year and two years later, the difference was about the same, although smoking increased during the follow-up. In 1986, six years after the program had terminated, an increase in smoking between ages 16 and 21 was seen (Figure 1). The differences between the direct program (38 percent smokers) and reference schools (42 percent smokers) were minimal, while the differences between the county-wide program and control schools were more impressive (31 vs 42 percent). When only those students who reported they were nonsmokers in the pretest were included in the analyses, the proportion of nonsmokers was lower in both the direct and teacher-led schools than in the controls at the eight-year follow-up (Table 1). If attrition is taken into account by assuming all dropouts to be smokers, smoking rates in the two treatment groups are more similar (48–49 percent) but they remain lower than the rate of smoking in the campaign group (59 percent). This adjustment probably overestimates smoking prevalence.

TABLE 1—Proportion of Smokers (any smoking) by School Group in the Surveys among Nonsmokers in the Beginning of the Program

				Difference (95% CI)	
	Direct Program (%) (N)	County-wide Program % (N)	Control Schools % (N)	Direct vs Control %	County-wide vs Control %
Pretest	0 (249)	0 (243)	0 (251)		
Posttest	19 (248)	19 (242)	29 (250)	10 (3,17)	10 (3,17)
Two-and-one-half-year follow-up	23 (213)	22 (216)	38 (231)	15 (7,23)	16 (8,24)
Four-year follow-up	28 (230)	26 (225)	41 (238)	13 (4,21)	15 (7,23)
Eight-year follow-up	37 (200)	31 (182)	47 (209)	10 (1,19)	16 (6,25)

CI = confidence interval

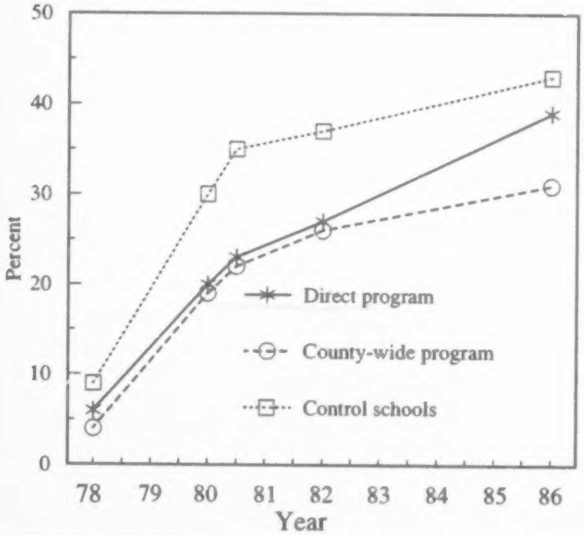

FIGURE 1—Proportion of Students Who Reported Smoking at least Once a Month in Different Surveys

Discussion

The interpretation of these results is complicated by several factors. Implementation evaluation of the project indicated that students at the urban school in the county-wide condition received only one or two anti-smoking sessions as part of the regular health education curriculum. The urban junior high school in the control condition shared its school yard with a senior high school whose students were allowed to smoke in the school yard. The primary difference that was observed, lower smoking rates, could be attributed to the overall community program in North Karelia rather than the specific programs that were carried out in the study schools. Given these qualifications, we cannot conclude from this study that preventive programs at school have long-term effects on smoking. On the other hand, our findings are consistent with the hypothesis that community and school programs can influence youthful nonsmokers to avoid tobacco.

ACKNOWLEDGMENTS

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The Effects of Detoxification of Domestic Gas on Suicide in the United States

David Lester, PhD

Abstract: As domestic gas was detoxified in the United States, the rate of suicide by domestic gas decreased. During this time period (1950–60), there was a parallel increase in the per capita ownership of cars and an accompanying increase in the rate of suicide by motor vehicle exhaust. However, displacement of suicide method from domestic gas to car exhaust occurred only for males and not for females. (*Am J Public Health* 1990; 80:80–81.)

Introduction

It has been argued that the detoxification of domestic gas in England and Wales in the 1960s and early 1970s was responsible not only for a drop in the rate of suicide by domestic gas, but also for a drop of about one-third in the overall suicide rate.[1,2]

However, it has also been noted that detoxification of domestic gas in Scotland and The Netherlands was not accompanied by a decrease in the overall suicide rate. Clarke and Mayhew[3] have suggested that this reflected a rising suicide rate in Scotland and The Netherlands at the time of detoxification which masked any reduction made possible by the detoxification of domestic gas.

Lester[4] has argued that, in the states of the USA where the availability of firearms is reduced by strict gun control laws, there is an accompanying reduction not only in the rate of suicide by firearms, but also in the overall suicide rate. Rich and Young[5] have shown that in Toronto people switched from using firearms to jumping in front of subway trains, leaving the suicide rate unchanged.

Domestic gas in the United States contained carbon monoxide in the early part of this century. The carbon monoxide content was reduced gradually during the 1940s and 1950s as gas companies switched to natural gas. The present paper examines the relation of the detoxification of domestic gas with rate of suicide.

During this same period, the per capita ownership of cars increased in the USA. The per capita ownership of cars in the states of America is positively associated with the use of car exhaust for suicide.[6] The present paper, therefore, also explored whether people used car exhaust rather than domestic gas as a suicide method as the relative availability of these two methods for suicide changed.

Method

Data on the rates of suicide by domestic gas and motor vehicle exhaust were obtained from the annual volumes, *Vital Statistics of the United States*, published by the National Center for Health Statistics. Data on the percentage of customers receiving natural gas were obtained from the annual volume *Statistical Abstract of the United States*, published by the Bureau of the Census. Data on the number of cars per capita each year were obtained from *Historical Statistics of the United States from Colonial Times to 1970* published by the Bureau of the Census.

Data were obtained each year from 1946, the first year for which data on domestic gas customers could be obtained using a consistent set of categories, to 1970, by which time almost all domestic gas customers were receiving natural gas. Suicide by natural gas is possible, but since natural gas is less toxic, suicide is by the exclusion of air containing oxygen and is less apt to occur.

Results

Table 1 shows that the percentage of domestic gas customers receiving natural gas increased from 1950 to 1970 (regression coefficient = 22.75, standard error = 1.86), as did the number of cars per capita (regression coefficient = 8.90, standard error = 0.22). The accidental death rate from domestic gas was strongly and positively related to the percentage of customers receiving the toxic manufactured or mixed gas and with the rate of suicide by domestic gas.*

The rate of suicide by domestic gas decreased over the years (regression coefficient = −34.70, standard error = 4.12), while the rate of suicide by motor vehicle exhaust increased (regression coefficient = 29.45, standard error = 2.52). The sum of these two suicide rates remained constant

TABLE 1—Suicide Rates 1950–70, using Domestic Gas and Motor Vehicle Exhaust and the Extent of Use of Natural Gas by Consumers

			Suicide Rates per 100,000 per Year by:				
	Cars per Capita	% of Customers Using Natural Gas	Domestic Gas	Motor Vehicle Exhaust	Total Gas/ Exhaust	All Other Methods	Overall
1950	0.26	60.8	0.73	0.60	1.33	9.93	11.26
1955	0.31	81.6	0.19	0.74	0.93	9.17	10.10
1960	0.34	92.8	0.10	0.91	1.01	9.53	10.54
1965	0.39	98.0	0.08	1.09	1.17	9.90	11.07
1970	0.43	98.6	0.02	1.02	1.05	10.40	11.45

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*Complete tables of data are available on request to author.

TABLE 2—Changes in the Suicide Rates 1950–70, by Method for Men and Women

	Suicide Rates by Men				Suicide Rates by Women			
	Motor Vehicle Exhaust	Domestic Gas	Total Gas	Other Methods	Motor Vehicle Exhaust	Domestic Gas	Total Gas	Other Methods
1950	1.09	0.85	1.94	15.75	0.12	0.61	0.73	4.32
1955	1.29	0.23	1.52	14.33	0.21	0.16	0.37	4.18
1960	1.55	0.13	1.68	14.78	0.29	0.07	0.36	4.59
1965	1.74	0.11	1.85	14.69	0.51	0.06	0.57	6.04
1970	1.48	0.03	1.51	15.30	0.61	0.01	0.62	5.95

during this period (regression coefficient = −3.81, standard error 6.67). The rates of suicide by domestic and motor vehicle gas were negatively related (r = −0.94). The total suicide rate during this period did not change systematically (regression coefficient = 1.84, standard error = 15.56).

The data were examined separately for men and women (Table 2). The use of domestic gas for suicide by men declined from 1950 to 1970 while the use of car exhaust increased. The use of both methods varied greatly from year to year, but the data suggest a slight decrease over the period. For women, the results are similar, except that the data suggest a drop in the use of domestic gas and car exhaust from 1950 to 1960 followed by a rise from 1960 to 1970.

Discussion

It has often been argued that removal of one lethal method for suicide will simply cause would-be suicides to turn to alternative methods.[7] The present results indicate that as domestic gas was detoxified and as per capita ownership of cars increased, fewer people used domestic gas for suicide while more used motor vehicle gas for suicide.

The situation is considerably more complicated, of course. During this period (1950 to 1970), the availability and use of firearms for suicide was increasing,[8] as perhaps was the availability of lethal medications such as antidepressants. Thus, as domestic gas became less toxic, several other methods for suicide became more available. This, together with the lesser use initially of domestic gas for suicide in the US as compared to England and Wales, is perhaps why detoxification of domestic gas in the US had little impact on the overall suicide rate, while in England and Wales detoxification of domestic gas was accompanied by a one-third reduction in the overall suicide rate.[7]

The data are not clear on whether switching occurred from domestic gas to car exhaust as the relative availabilities of these methods changed. A case could be made for or against switching, depending upon which time period (1950–60 or 1960–70) and which sex is examined. Two important questions for future research are: "Which people will switch methods for suicide if one is made less available?" and "To which methods will people switch if one method is made less available for suicide?"

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I should like to thank Dr. Ronald Clarke for advice in the preparation of this paper.

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NIA Funds Three New Alzheimer's Disease Research Centers

The National Institute of Aging (NIA) has announced funding of three new Alzheimer's Disease Research Centers (ADRCs), with $2.3 million for the first year of operation and a total of $11 million over the next five years. This brings to 15 the number of Alzheimer's centers funded by the NIA. The purpose of the ADRC program is to enhance research on Alzheimer's disease by providing the resources and environment for collaborative, multidisciplinary studies. Each center will serve as the site for new and expanded studies of the basic, clinical and behavioral aspects of Alzheimer's disease. In addition to their research mandate, the centers train scientists and health care professionals and offer consultation and education to Alzheimer patients and their families.

The three new centers are located at Baylor College of Medicine, Houston, Texas; Columbia University, New York City; and the University of Michigan, Ann Arbor.

The other 12 NIA-funded centers are located at: Duke University, Durham, NC; Harvard Medical School/Massachusetts General Hospital, Boston, MA; Johns Hopkins Medical Institutions, Baltimore, MD; Mt. Sinai School of Medicine/Bronx VA Medical Center, New York; University of California-San Diego, La Jolla, CA; University of Kentucky, Lexington, KY; University of Pittsburgh, Pittsburgh, PA; University of Southern California, Los Angeles; Case Western Reserve University, Cleveland, OH; University of Washington, Seattle, WA; Washington University, St. Louis, MO; and University of Texas, Southwestern Medical Center, Dallas, TX.

Variables Influencing Condom Use among Intravenous Drug Users

STEPHEN MAGURA, PHD, JANET L. SHAPIRO, MA, QUDSIA SIDDIQI, PHD, AND DOUGLAS S. LIPTON, PHD

Abstract: Correlates of condom use were identified using cross-sectional data from a convenience sample of 211 sexually active intravenous drug users enrolled in methadone maintenance in New York City. Sixty-eight percent did not use condoms at all in the previous month and only 11 percent used condoms every time. Nineteen percent were planning on conception, only 20 percent of whom had been tested for human immunodeficiency virus (HIV) antibody (all seronegative). Multiple logistic regression analysis indicated that condom use was independently associated with greater personal acceptance of condoms, greater partner receptivity to sexual protection, and recent entry to methadone treatment. *Am J Public Health* 1990; 80:82–84.)

Introduction

Thirty percent of adult AIDS (acquired immunodeficiency syndrome) cases diagnosed in the United States during 1988 involved intravenous drug users (IVDUs).[1] In New York City, at least 50 percent of IVDUs are already HIV (human immunodeficiency virus) seropositive.[2,3] There has been extensive public health focus on reducing needle-sharing and other drug-use-related HIV transmission risks within the IVDU population.[4] However, sexual contact is an important secondary means of HIV transmission among IVDUs and is the main route of HIV transmission to the general heterosexual population.[5,6] Heterosexual contact accounted for 5 percent of US adult AIDS cases diagnosed during 1988.[1] Infected IVDU women are also the primary source of perinatal HIV transmission.[7]

Condoms are an accepted measure to prevent the sexual transmission of HIV; they also can be a useful although not fully effective birth control method for seropositive women.[8–10] One recent study found that only 13 percent of IVDUs used condoms (usually inconsistently) and that 65 percent of their sexual contacts were with non-IVDUs; the sample was drawn from a treatment population with 57 percent HIV seroprevalence.[11] Overall there has been little or no change in sexual practices in response to the AIDS threat either among IVDUs[12–14] or the general heterosexual population.[6] We present a quantitative analysis of the factors influencing IVDUs' decisions about condom use.

Methods

The study sample is 289 IVDUs enrolled in methadone maintenance who had volunteered for an AIDS prevention demonstration/research project at three New York City clinics. Because participants were self-selected, they may not be entirely representative of the clinic populations. This analysis involves the 211 subjects who had reported sexual activity within the previous month. The data are derived from baseline (pre-intervention) questionnaires completed between January 1987 and February 1988. The self-administered questionnaire consisted of fixed-response, Likert-type items and was available in both English and Spanish. Research staff were present on site to assist subjects who needed help in completing the instrument. Confidentiality was guaranteed; the results were not shared with methadone clinic staff. Subjects were paid $20 in two of the clinics and $10 in the third for completing the questionnaire. The questionnaire included measures of knowledge about AIDS and its transmission; drug use and needle sharing; sexual activities and condom use; and attitudes and beliefs related to condom use.

The dependent variable was condom use, measured by the item: "In the last month, how often did you (or your partner) use a rubber (condom) for sex?" ("Never," "sometimes," "most times," or "every time"). Responses to this question were reasonably internally consistent; subjects reporting at least some use usually disagreed with a different, negatively phrased item on condom use: "I almost never have sex using rubbers (condoms)" ($r = -.70$). For the analysis condom use was dichotomized as no use vs some use in the last month. The independent variables were drawn from theory and research on the determinants of condom use in the general population[6] and among gay/bisexual men.[15]

Results

Sixty-eight percent of the subjects did not use condoms at all in the previous month and only 11 percent used condoms every time for sex. Table 1 presents a multiple logistic regression on condom use. Variables independently related to condom use were: greater personal acceptance of condoms, greater partner receptivity to sexual protection, and recent entry to methadone treatment.

Discussion

The low rate of condom use among these subjects is consistent with findings from other surveys of addicts,[11,16] gay/bisexual men,[15] and the general population.[6]

The multivariate results indicate that norms and expectations within sexual partnerships exert the primary influences on condom use for these IVDUs. Condom acceptability (believing that condoms do not impair enjoyment of sexual relations and being willing to use condoms when requested by a sexual partner) was the variable most strongly associated with condom use. Perceived partner receptivity to suggestions of condom use and other sexual risk reduction was also directly related to condom use. That is, when individuals are amenable to condom use, and when sexual partners effectively communicate reciprocal willingness, there is greater likelihood of condoms being used in that relationship.

Beliefs, attitudes, and communication skills involving condoms can be addressed by methods such as cognitive-behavioral group work with addicts and, when possible, with their sexual partners.[17] In groups participants who already know that condoms do not interfere with sexual pleasure can encourage or challenge others to verify this for themselves; experience with condoms is likely to decrease resistance to

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From Narcotic and Drug Research, Inc. Address reprint requests to Stephen Magura, PhD, Deputy Director of Research, Narcotic and Drug Research, Inc., 11 Beach Street, New York, NY 10013. Ms. Shapiro is a Research Associate, Dr. Siddiqi is an Assistant Project Director, and Dr. Lipton is Director of Research at NDRI. This paper, submitted to the Journal February 27, 1989, was revised and accepted for publication June 26, 1989.

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TABLE 1—Multiple Logistic Regression on Condom Use

Variables	(N)	Percent Condom Use	Odds Ratio (95% Confidence Intervals)
Gender			
Male	(132)	31	1.00
Female	(79)	33	1.16 (.74, 1.84)
Ethnicity			
White[a]	(144)	31	1.00
Black	(31)	39	1.47 (.85, 2.55)
Hispanic	(36)	31	.96 (.54, 1.69)
Age (years)[b]			1.30 (.91, 1.84)
Under 30	(45)	31	
30–34	(76)	33	
35–39	(58)	28	
40 & over	(30)	33	
Marital status			(Not entered)[c]
Single, never married	(85)	39	
Married	(77)	26	
Other	(49)	29	
Living with a Sexual Partner[d]			
No	(74)	43	1.00
Yes	(137)	25	.70 (.44, 1.12)
Sexual Partners, Past Year			
One	(144)	27	1.00
Two or more	(66)	42	1.06 (.65, 1.71)
Sexual Partners Injecting Drugs, Past Year			
None	(140)	29	1.00
One or more	(70)	37	.98 (.62, 1.55)
Planning on Conception			
No	(170)	35	1.00
Yes	(39)	18	.84 (.46, 1.50)
Frequency of Heterosexual Relations, Past Month			1.91 (.83, 4.39)
One or two times	(84)	26	
One day a week	(26)	38	
A couple days a week	(78)	38	
Almost every day	(20)	20	
IV Drug Use Risk, Past Month[e]			
None	(125)	30	1.00
Injects, no needle sharing	(50)	34	1.09 (.67, 1.76)
Injects and shares needles	(36)	36	1.11 (.62, 1.97)
Knows AIDS/ARC Victims[f]			
None	(130)	26	1.00
One or more	(81)	41	1.50 (.98, 2.29)
Knows HIV Serostatus			
No	(174)	30	1.00
Yes, Seronegative	(30)	40	.80 (.29, 2.21)
Yes, Seropositive	(7)	43	1.39 (.81, 2.39)
Fatalism about AIDS[g]			
Agrees will get AIDS	(48)	40	1.00
Disagrees will get AIDS	(160)	29	.79 (.60, 1.62)
Condom Reliability[h]			
Agrees condoms break	(124)	27	1.00
Disagrees condoms break	(82)	40	1.12 (.73, 1.70)
Knowledge of Sexual Risks[i]			.91 (.71, 1.16)
Low	(52)	36	
Medium	(98)	34	
High	(61)	24	
Condom Acceptability[j]			1.75 (1.21, 2.52)
Low	(29)	10	
Medium	(116)	24	
High	(63)	56	
Condom Efficacy against AIDS[k]			1.13 (.76, 1.67)
Low	(91)	24	
Medium	(53)	28	
High	(66)	44	
Partners' Receptivity to Protection[l]			1.47 (1.02, 2.12)
Low	(24)	13	
Medium	(47)	26	
High	(135)	37	
Perceived Past Exposure to HIV[m]			1.07 (.75, 1.52)
Low	(33)	21	
Medium	(75)	35	
High	(103)	33	
Perceived Susceptibility to AIDS[n]			1.27 (.90, 1.79)
Low	(55)	27	
Medium	(88)	26	
High	(67)	42	
Time in Methadone Program[o]			.83 (.72, .96)
1–11 months	(50)	40	
12–23 months	(43)	39	
24–35 months	(27)	26	
36–47 months	(11)	0	
48 months and over	(78)	28	

continued

TABLE 1—Continued

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[a] Includes 5 "others."

[b] Entered in regression as a continuous variable. Odds ratio is for a 5-year age interval.

[c] Not entered in the logistic regression because of high association with next variable, "Living with a Sexual Partner."

[d] "Are you now living with someone with whom you have a sexual relationship?"

[e] "In the last month, how often did you shoot up any drugs?" and "In the last month, how often did you share a needle or works with someone?"

[f] "How many people with AIDS or ARC (AIDS-related complex) do you know personally?"

[g] "My gut feeling is, I'm going to get AIDS sooner or later."

[h] "I don't like using rubbers (condoms) because they sometimes break."

[i] The correct answers were summed for seven items: "You could get AIDS by having unprotected sex with a person who shot drugs" (T); "only homosexuals can give you AIDS through sexual relations" (F); "pregnant women who carry the AIDS virus can pass it to their newborn children" (T); "anal intercourse is the only way sex can give you AIDS (F); "a positive AIDS antibody test means you could give the AIDS virus to someone through sex" (T); "you can give the AIDS virus to someone else by sex or needle-sharing even if you feel healthy" (T); "you can't get AIDS just by oral sex (F)." Higher knowledge = higher score. Entered in regression as a continuous variable.

[j] Two items were summed: "I don't like using rubbers (condoms) because they cut down on my enjoyment" (agree = 0, not sure = 1, disagree = 2); "I would use rubbers (condoms) if my partner asked me" (agree = 2, not sure = 1, disagree = 0). Higher acceptability = higher score. Entered in regression as a continuous variable.

[k] Two items were summed: "Using a rubber (condom) can help protect you from getting AIDS" (agree = 2, not sure = 1, disagree = 0); "there's no way to have 'safe sex' with someone who's carrying the AIDS virus" (agree = 0, not sure = 1, disagree = 2). Higher perceived efficacy = higher score. Entered in regression as a continuous variable.

[l] Two items were summed: "My sex partner would be insulted if I suggested that we use a condom" (agree = 0, not sure = 1, disagree = 2); "my sex partner would be insulted if I suggested any changes in the way we have sex" (agree = 0, not sure = 1, disagree = 2). Greater receptivity = higher score. Entered in regression as a continuous variable.

[m] Two items were summed: "I've already done plenty that could have exposed me to AIDS" (agree = 2, not sure = 1, disagree = 0); "I never did anything that could give me AIDS" (agree = 0, not sure = 1, disagree = 2). More exposure = higher score. Entered in regression as a continuous variable.

[n] Two items were summed: "I'm worried that some of my sexual partners could give me AIDS" (agree = 2, not sure = 1, disagree = 0); "I'm worried about my chances of getting AIDS" (agree = 2, not sure = 1, disagree = 0). Higher susceptibility = higher score. Entered in regression as a continuous variable.

[o] Odds ratio is for a 6-month interval. Entered in regression as a continuous variable.

their use.[18] Participants can also practice communication and negotiation skills pertinent to sexual relationships. Even if neither partner personally objects to condom use, as our data suggest is often true, one partner must nevertheless take the initiative and broach the subject of use.

Addicts who recently entered methadone treatment were more likely to use condoms. This relationship requires more detailed investigation.

It is of particular concern that, of the 39 subjects (19 percent) planning on conception, only 20 percent reported being tested for HIV antibody (all seronegative). These findings held true equally for men and women. AIDS health education must emphasize the importance of HIV testing for IVDUs who wish to conceive, given the at least 30 percent risk that seropositive women will bear infants who will develop HIV-related illness.[19,20]

The study has several limitations. Causal interpretations should be considered tentative because the data are cross-sectional. The subjects were all in methadone treatment; an estimated 15 percent of IVDUs in New York City are in treatment at any given time.[21] It is possible that a study of street addicts would yield different results, but note that we found no association between current intravenous drug use and condom use for our sample (Table 1). Not all variables hypothesized or found to be associated with condom use in the research literature were measured by the study. These include a possible belief that condoms make sex seem premeditated, possible embarrassment in purchasing condoms, and the cost of condoms for low-income persons. The first two factors, however, have seemed particularly important only for adolescents.[22–24] The cost of condoms has become a less pertinent consideration for addicts since many drug abuse treatment programs (including the study clinics), as well as AIDS prevention outreach projects, have been distributing condoms for free.

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1990 SEPTEMBER 30–OCTOBER 4 NEW YORK, NEW YORK
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Results of an Intervention to Improve Compliance with Referrals for Evaluation of Suspected Malignancies at Neighborhood Public Health Centers

CLARA MANFREDI, PHD, LORETTA LACEY, RN, DRPH, AND RICHARD WARNECKE, PHD

Abstract: A study was conducted at Neighborhood Health Centers to evaluate procedures to improve compliance with referrals of patient at risk for cancer. Intervention consisted of a standardized communication from the exit nurse, a patient form to be returned after compliance, and one written and one telephone reminder as needed. Compliance was 68.2 percent in a control group and 89.0 percent among patients who received the experimental intervention, a significant increase of 20.8 percent (95% CI: 12.5, 29.1). (*Am J Pubic Health* 1990; 80:85–87.)

Introduction

Excessive cancer mortality among low income and minority populations could be reduced with better screening and earlier detection and treatment.[1–6]

Indigent populations rely upon public clinics for primary care which refer persons with suspected cancer to other medical facilities for further care; compliance with referrals is often problematic, contributing to delayed treatment.[7,8]

Delay in seeking care has been associated with difficulty communicating with physicians,[7,9,10] difficulty obtaining prompt appointments,[7,11] and concern about financial costs.[12–15] In contrast, compliance improved with prompt appointment dates, clear and specific instructions, and reminders.[7,11,16–18] Fear of cancer may be a deterrent to promptly seeking a diagnosis,[19,20] whereas knowledge of cancer symptoms and belief in the possibility of early detection and cure may lead to prompt diagnosis and treatment.[13,14,20–22] This study, however, focused only on intervention designed to improve referral information and remind those patients who may not have kept the referral appointment.

Methods

The study was conducted between October 1984 and January 1986 at four Chicago Department of Health Neighborhood Health Centers (NHCs). Subjects were adult medicine and maternal and family planning patients, at least 18 years old, without prior cancer history and referred for treatment or further diagnostic workup for a suspected cancer.

At the exit interview, nurses identified eligible patients and, in the experimental condition, administered the intervention. Overall, 292 control and 226 experimental group patients were accrued. No patient consent was needed because the intervention was an improvement on normal care and placed the subject at no risk. Monitoring procedures revealed that approximately 50 percent of eligible cases were not included in the study because of direct physician referral, missed opportunity by nurses to see the patient, nurse was too busy. Missed cases were more likely than those accrued to the study to be referrals for non-specific evaluation and to lack an appointment date, but occurred equally in the control and experimental conditions based on NHC assignment. Study results are based only on identified cases.

Unpredictability of case occurrence and need to integrate the study procedures into routine NHC's activities made random assignment to study conditions unfeasible. Instead, each NHC was assigned to experiment or control on alternate months, with two NHCs in either condition in any given month.

The experimental intervention consisted of:

- A verbal nurse-patient communication, structured with a mini-questionnaire designed to facilitate queries about referral procedures (i.e., why needed, where to go, whom to contact, etc.).
- A form, which patients were requested to return in a prestamped, pre-addressed envelope after referral completion.
- A reminder note, mailed to 163 patients (72.1 percent) whose chart had no return form nor other record of compliance two weeks after appointment or six weeks after referrals without appointment.
- A telephone reminder, attempted with 102 patients (45.1 percent) still without record of compliance three weeks later. Only 79 calls were completed, because patients could not be found or nurses could not make numerous attempts. Those not contacted by phone were sent a second written reminder and 13 patients were reminded of the referral as they revisited the NHC.

Control patients received the NHCs' usual referral procedures.

Compliance was ascertained by phone interviews. Interviews were attempted with all patients three weeks after final chart review and completed with 393. The remaining were: 106 not located, 15 with communication barriers, two deceased, and two refused. Completion was 75.3 percent in the control and 76.5 percent in the experimental group. Study results are based only on the 393 cases with final interview.

Results

Compliance was reported by 304 patients (77.5 percent). Of these, 61 (27.7 percent) of the control group and 73 (42.2 percent) of the experimental group also had documentation of compliance in their medical chart and 51 patients (29.5 percent of experimental group) had returned their patient intervention form.

Study patients were predominantly Black (69.5 percent), female (88.5 percent), 48 percent were under age 55, and 23 percent under age 35. Only 18.8 percent of referrals received

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TABLE 1—Percent of Patients Who Complied with Referrals in Each Study Group, Overall and by Patient and Referral Characteristics

	Control Group (N = 220)		Experimental Group (N = 173)			
Variables	%	N	%	N	Per cent Difference	(95% CI)
Overall	68.2	(150)	89.0	(154)	20.8	(12.5, 29.1)
Sex						
Male	84.6	(22)	100.0	(19)	15.4	(−1.6, 32.4)
Female	66.0	(128)	87.7	(135)	21.7	(12.7, 30.7)
Race						
Black	64.7	(101)	87.2	(102)	22.5	(12.5, 32.5)
White	61.1	(11)	77.8	(7)	16.7	(−16.3, 49.7)
Hispanic	88.9	(8)	96.0	(24)	7.1	(−23.9, 38.1)
Others	81.5	(22)	94.7	(18)	13.2	(−6.5, 32.9)
Age (years)						
18 to 34	73.7	(28)	90.3	(23)	16.6	(−2.0, 35.2)
35 to 44	65.0	(13)	100.0	(25)	35.0	(13.4, 56.6)
45 to 54	69.6	(16)	79.3	(23)	9.7	(−14.0, 33.4)
55 to 64	63.8	(30)	87.5	(36)	23.7	(5.5, 41.9)
65 to 74	65.4	(34)	96.3	(26)	30.9	(11.0, 50.8)
75 or more	72.5	(29)	81.0	(17)	8.5	(−14.4, 31.4)
Education (years)						
Less than 9	71.6	(58)	92.2	(47)	20.6	(6.5, 34.7)
9 to 11	68.1	(49)	86.2	(150)	18.1	(3.4, 32.8)
12 to 14	67.2	(39)	89.5	(51)	22.3	(7.2, 37.4)
15 or more	44.4	(4)	85.7	(6)	41.3	(−6.5, 89.1)
Clinical Reasons for Referral						
Prior test results	81.6	(40)	97.1	(33)	15.5	(1.3, 29.7)
Symptoms only or missing information	64.3	(110)	87.0	(121)	22.7	(12.9, 32.5)
Procedure Requested in the Referral						
Evaluate, rule out	83.3	(50)	93.0	(53)	9.7	(−2.1, 21.5)
Mammography	54.7	(52)	84.1	(58)	29.4	(14.8, 44.0)
All others	73.8	(48)	91.5	(43)	17.7	(3.0, 32.4)
Appointments						
Within 15 days	92.9	(39)	96.9	(31)	4.0	(−6.4, 14.4)
Within 16 to 147 days	53.5	(38)	82.5	(52)	29.0	(13.1, 44.9)
No appointment	68.2	(73)	91.0	(71)	22.8	(10.7, 34.9)
Insurance Coverage						
Yes	67.5	(77)	89.6	(60)	22.1	(9.2, 35.0)
No	68.9	(73)	88.7	(94)	19.8	(7.9, 31.7)
Referral Hospital						
Cook County Hospital	60.8	(79)	86.8	(105)	26.0	(15.0, 37.0)
Other hospitals	78.9	(71)	94.2	(49)	15.3	(2.9, 27.7)

an appointment date within 15 days and 47.1 percent received no appointment. Almost half of the patients had no insurance and 63.9 percent were referred to the public county hospital.

Experimental group patients were more likely than controls to be Hispanic (14.7 versus 4.1 percent), lack insurance (61.3 versus 48.2 percent) and be referred to the county hospital (69.9 versus 59.1 percent).* Within the experimental group, patients who had received an appointment date were more likely to have returned their intervention form (36.8 percent) than those without an appointment date (20.5 percent).

Compliance with referrals was 68.2 percent in the control and 89.0 percent in the experimental group, an increase of 20.8 percent (Table 1). Characteristics associated with better compliance in both groups included: male gender, referrals following prior test results and for appointments within 15 days. Compliance was less likely for referrals for Blacks, for mammography, and for referrals to the county hospital.

Improved compliance was most evident for conditions associated with non-compliance in the control group. It was least evident for appointments within 15 days, which were associated with high compliance regardless of intervention.

The high appointment-keeping of patients with prompt appointments is striking and may be due partly to the latter being associated with more urgent medical conditions. Appointments within 15 days were more likely to be for follow-up of positive Pap smears (47.4 percent) than for positive guaiac or other laboratory results (28.9 percent) or for all other conditions (13.9 percent) and for referrals for surgical or other procedures (33.7 percent) than for evaluation (20.5 percent) or mammography (9.1 percent) (data not presented in tables).

Because the majority of patients in the study were Black females, results were analyzed separately for this group. Compliance among Black females was 61.9 percent in the control and 85.8 percent in the experimental group. Their patterns of compliance by patient and referral variables within study group were very similar to those reported in Table 1 for the entire sample.

Discussion

The study findings indicate that compliance with referrals can be improved when appointments are within two weeks of referral and when clinic staff give clear instructions and follow-up patients to ensure receipt of prescribed procedures. Even when the situation of the referral would predict poor compliance, careful attention by the clinic staff reduced the noncompliance significantly.

The majority of referrals in this study were for mammography and to follow-up positive Pap smears. Without the intervention, compliance with referrals following prior test results (thus including positive Pap smears), was 81.6 percent, but only about half of the women referred for mammography received the procedure.

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*Exact data available on request to authors.

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APHA Invites Proposed Policy Statements for 1990 Consideration

The American Public Health Association has announced the schedule for the annual process of considering proposed policy statements. Individual members of APHA may submit resolutions. The various sections, affiliates, caucuses, SPIGs (special primary interest groups), and other components defined in the APHA constitution and by-laws may offer position papers.

Resolutions are short, single statements of 500 words or less, focusing on a specific action or series of actions endorsed by the Association, whereas a *position paper* is a major exposition, of up to 3,000 words, of the Association's viewpoint on a broader issue affecting the public's health. Each proposed statement will be considered by the APHA Joint Policy Committee which may recommend revisions to authors before the time of the APHA annual meeting. When the convention opens in New York City, open public hearings are held on all proposed policy, and the Governing Council takes final action by adopting or rejecting the proposed policy statements.

These public policy statements form the basis of APHA's stand on legislative, regulatory, legal, and health matters of concern to the public health community. They may stimulate scientific inquiry or be intended to initiate government action. The statements also provide a record of APHA's position on a variety of health issues, reflecting the diverse interest of the Association's membership, and further serve as a chronicle of changing and evolving public health concerns over the years of American history.

The *deadline* for receipt of proposed resolutions and position papers for consideration this year is *February 26, 1990*. A copy of the schedule and procedures is available from: Joint Policy Committee, APHA, 1015 15th Street, NW, Washington, DC 20005.

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Fecal Contamination of Shanty Town Toddlers in Households with Non-corralled Poultry, Lima, Peru

Grace S. Marquis, MS, Gladys Ventura, BS, Robert H. Gilman, MD, Esperanza Porras, BS,
Elba Miranda, MS, Luz Carbajal, MS, and Marta Pentafiel, BS

Abstract: We used direct observer techniques to measure the frequency with which toddler-aged children were contaminated by poultry feces in homes in a peri-urban shanty town in Lima, Peru. The mean number of fowl was 5.4 (SD 3.1), with 10.0 (SD 10.7) poultry defecations per 12 hours. Toddlers' hand contact with poultry feces occurred a mean of 2.9 (SD 3.0) times/12 hours. A mean of 3.9 (SD 4.6) feces-to-mouth episodes per household/12 hours occurred both by direct hand-to-mouth contamination and indirectly by handling soiled objects which were then placed in the mouth. There was a strong correlation between feces-to-hand contamination and feces-to-mouth contamination ($r = 0.94$). There was also an association between feces-to-mouth contamination and the number of stools deposited in the house ($r = 0.66$). For each additional chicken stool deposited during the day, there was an average increase of 0.27 in feces-to-mouth episodes.

We collected feces from 68 infected chickens and found viable *Campylobacter jejuni* for up to 48 hours after deposition. Yet, a survey of 108 families demonstrated that free-roaming poultry were often not thought of as a health risk for children. An intervention program to reduce oral-fecal contamination should emphasize that all poultry be corralled and not allowed access into the house. (*Am J Public Health* 1990;80:146–149.)

Introduction

Diarrheal disease is an important factor in the infection-malnutrition cycle of infants and young children in developing countries. In Peru, children under three years of age have a diarrheal incidence of 10 episodes per child-year, which can significantly affect their growth rates.[1] One important cause of diarrhea in Lima children less than one year of age is *Campylobacter jejuni*. This organism was found in 15 percent of diarrhea stool samples[2] and accounted for the majority of dysenteric diarrheas.[3]

In Peru, *C. jejuni* has been isolated from 88 percent of commercial chickens and 50 percent of free-ranging domestic chickens in a pueblo joven (a peri-urban shanty town community).[4] In an aged-matched case-control study of Lima peri-urban households, children in families with chickens had a three-fold relative risk of having *C. jejuni* diarrhea which increased to 12.5 in families with *C. jejuni* infected chickens.[2] These risk factors suggest that direct contact with free-roaming *C. jejuni*-infected chickens that defecate on household surfaces is an important transmission vehicle for young children.

Direct observation of the personal hygiene of small children and mother-child behavioral patterns is the most reliable method for obtaining data.[5] We used directed observer techniques to measure the rate at which toddlers in a shanty town contaminate themselves with domestic poultry feces.

Methods

Study Sites

The study was done in two pueblos jovenes (recent shanty towns): Huascar, Canto Grande[6] and Las Pampas de San Juan, San Juan de Miraflores, in Lima Peru. The majority of the study population are rural immigrants who are daily laborers or street vendors with low, unstable incomes. Housing materials in Huascar and San Juan range from woven cane matting to permanent structures of brick and cement. Although the majority of pueblos jovenes do not have electrical installations, most of our study households had them. Water is scarce in both pueblos. It is provided through tanker trucks, and in half of San Juan through piped communal water spigot. In-house sewage systems are not available.

Direct Observation Data

Ten families from Huascar participated in an in-depth behavioral observation study, with a total of 21 children under five years of age. All the families had at least one child under two years of age, and domestic chickens which were not corralled at least part of the day. They gave informed verbal consent prior to their participation. The study received ethical approval by all involved Peruvian institutions (Universidad Peruana Cayetano Heredia, Instituto de Investigación Nutricional, and PRISMA - Proyectos en Informática, Salud, Medicina y Agricultura).

To gather the direct observation data, a field worker remained in the toddler's home for one 12-hour observation period (7am to 7pm). At the beginning of the visit, a descriptive inventory of the domestic animals was completed. During the 12-hour period the observer noted all activities in the house for each fowl, including the number of times the fowl entered the home, the duration of time spent inside, and how frequently and where the poultry defecated. The observer also tabulated the activities of the toddler, specifically noting the number of times the child's hands were contaminated with poultry feces (defined as feces-to-hand episodes) and the number of times that his/her hands, or other objects contaminated with poultry feces, were introduced into his mouth (defined as feces-to-mouth episodes). Feces-to-mouth episodes were considered to have been interrupted and were not tabulated if the child's contaminated hands had been washed before he placed them in his mouth.

All feces-to-hand and feces-to-mouth episodes were tabulated for each family and analyzed as episodes per child. Correlations and linear regressions for the feces-to-hand and

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feces-to-mouth episodes were performed with the following variables: number of fowl, hours in the house/12 hours, and number poultry stools/12 hours.

Poultry Raising Practices and Beliefs

We surveyed a group of 108 families living in Las Pampas de San Juan to document their practices and beliefs regarding the raising of chickens and other fowl. The survey identified the reasons families gave for raising or not raising poultry, their beliefs about corralling the poultry, and the effect that confinement has on the growth and the health of the animal, the perceived advantages of raising chickens in the home compared to buying them commercially, and the mothers' perception concerning whether poultry roaming within the house presented a health risk to young children.

Socioeconomic Data

As part of a longitudinal study of *Enterobius vermicularis*, 121 randomly selected families from Huascar were included in a socioeconomic survey. The same data were collected for the Pampas families that participated in the poultry raising practices survey. The socioeconomic survey detailed age, sex, level of education, and occupation of all members of the family. Questions on housing quality included the materials used to build the house, number of rooms, ownership of functioning radios, televisions, and refrigerators, and presence or absence of a latrine. Finally the number and types of animals belonging to the home owners were also registered.

The practices and beliefs of mothers who did and did not raise poultry were compared using Chi square analysis. Each socioeconomic data set from the two pueblos jovenes was analyzed individually, and compared for socioeconomic indicators, using Student t test and Chi-Square analysis.

Survivability of *C. jejuni* in Deposited Chicken Feces

Stool samples were taken from *C. jejuni*-positive chickens, raised in a household in Huascar, Canto Grande. A cloacal swab was taken immediately after the chicken defecated and another swab was taken from the stool on the ground. These served as baseline samples. The stool sample, on the ground of a sun-exposed patio, was fenced in with a metallic wire and marked with its respective identification number and code. After 24 and 48 hours, additional samples were taken. All the samples were immediately placed in Cary Blair media containing Skirrow's antibiotic supplement,[7] and processed by direct plating on blood agar with Butzler's supplement.[8] Another sample was inoculated into thioglycolate broth supplemented with iron and the following antibiotics: vancomycin 10 μg/ml, polymyxin B 2.5 I.U./ml and trimethoprim lactate 5 μg/ml. The plates were incubated immediately while the broth was refrigerated for 24 hours and then plated with Butzler's agar. All plates were incubated at 37°C under microaerophilic conditions. Campylobacter species were identified based on colony morphology, oxidase production, and spiral appearance in the Gram stain.

Results

Direct Observation in the Homes

The interaction between toddlers and poultry was observed for a total of 119 hours in 10 Huascar households. The majority of the poultry were chickens (fowl/house = 5.4, chickens/house = 4.4). Three out of 10 families also had ducks and turkeys. Chickens and ducks only were allowed to roam freely through the house for about one-third of the 12-hour observation period (Table 1). The number of poultry and the time they were in the house were correlated (r = 0.64), as were the number of observed stools and the time the animals spent within the house (r = 0.86).

The children had approximately three feces-to-hand episodes per 12 hours (Table 1). The risk of feces-to-hand contamination was from both fresh stool and from that deposited several days previously but not removed from the household surfaces. In three households where no observed defecations occurred, hand and oral contamination still happened with stools that had been deposited prior to the day of observation. The two most important variables were number of droppings and number of minutes chickens were in the house.

There were about four feces-to-mouth episodes per child (Table 1). Children only rarely washed their hands after touching feces, and often put their fingers in their mouths, so feces-to-hand and feces-to-mouth episodes were highly correlated (r = 0.94); feces-to-hand episodes and the number of poultry accounted for 93 percent of the variation in the dependent variable—the number of feces-to-mouth episodes. Neither feces-to-hand nor feces-to-mouth episodes were related to the number of young children in study families. For each additional chicken stool deposited during the day, there was an average increase of 0.27 in feces-to-mouth episodes.

Poultry Raising Practices

One hundred and eight mothers from the Pampas de San Juan participated in the poultry raising practices survey. Eighty-one (75 percent) of the mothers were migrants, usually from the sierra zones. Whereas 88 percent (14/16) of the Lima-born mothers raised multiple types of poultry, only half of the migrant mothers raised other types of poultry in addition to chickens (27/54).

The major reasons given for a raising poultry (Table 2) included their use for home consumption and enjoyment of raising animals. The majority of mothers realized that commercial chickens were cheaper than those raised at home. When mothers were asked whether commercial or domestic chickens cost more, 77/108 (71 percent) thought the commer-

TABLE 1—Number of Domestic Poultry and Their Interactions with Toddler-Aged Children, Direct Observation Households (n=10)

			Feces-to-Hand*		Feces-to-Mouth*	
	Mean	SD	Regr Coef	95 CI	Regr Coef	95 CI
# Foul per house	5.4	3.1	−0.0326	(−0.693+0.628)	0.2831	(−0.735+1.301)
# Min/12 hours fowl in house	14043.0	2099.8	0.0004	(−0.001+0.001)	0.0006	(−0.001+0.002)
# Stools/12 hours	10.0	10.7	0.1664	(−0.025+0.308)	0.2718	(+0.057+0.487)
# Feces-to-hand episodes	2.9	3.0			1.474	(+1.101+1.846)
# Feces-to-mouth episodes	3.9	4.6				

*Dependent variable in simple linear regression model.

TABLE 2—Pampas Family Practices and Beliefs Concerning the Raising of Domestic Poultry

Family Practices, Beliefs	Families with Poultry n	Families with Poultry % of total	Families without Poultry n	Families without Poultry % of total
Why they raise poultry:				
To have to eat	41	60		
Enjoy raising animals	17	24		
No response	11	16		
Why they do not raise poultry:				
Are filthy/carry disease			11	29
Not economic (more expensive/ are stolen)			8	21
Without infrastructure/time			12	32
Do not like it/not accustomed			7	18
Where poultry grow best:				
Corralled	27	39	22	58
Roaming freely	37	54	13	34
Either or both places	4	6	2	5
No response	1	1	1	3
Is there a health risk for the family, having poultry loose in the house?				
No	20	29	11	28
Yes, cause illness	16	23	15	38
Yes, dirty or bring flies	13	19	8	21
Don't know	20	29		13

TABLE 3—Huascar and Pampas Family Socioeconomic Status, According to Selected Variables

	Huascar families		Pampas families	
Mean family size	6.5 S.D. 2.5		5.1 S.D. 1.8	
Mean age of mother(yrs)	31.5 S.D. 9.8		27.1 S.D. 5.8	
Mean number of years in Lima	—		15.8 S.D. 7.8	
	n	% of total	n	% of total
% with radio	79/120	66	77/108	71
% with television	82/120	68	73/108	68
% with radio and/or television	108/120	90	93/108	86
% with refrigerator	22/120	18	19/108	18
% with bore hole latrine	80/120	67	73/108	68
Housing quality:				
% with noble materials (bricks)	42/120	35	18/108	17
Education of mother:				
% completing secondary	15/120	13	31/108	29
% completing primary	60/120	50	56/108	52
% not completing primary/none	43/120	36	21/108	19
% no response				
Education of father:	2/120	2		
% completing secondary	31/120	26	56/108	52
% completing primary	66/120	55	42/108	39
% not completing primary/none	14/120	12	3/108	3
% absent	5/120	4	4/108	4
% no response	4/120	3	3/108	3

cial chickens were cheaper while only 22/108 (20 percent) thought the reverse.

In 83 percent (58/70) of the families, the poultry roamed freely for much of the day, with access into the house. This practice was explained in part by the belief of more than one-half of the mothers that the animals grew better when they were allowed to wander freely. Seventy four percent of the mothers not raising poultry and 68 percent of the mothers with poultry thought that there was a health risk or a negative hygienic aspect associated with free-roaming poultry.

Socioeconomic Data

Table 3 shows the socioeconomic indicators studied in both pueblo joven communities. No socioeconomic differences were found in families that did or did not raise chickens and other poultry.

Microbiological Results

Forty percent (27/68) of the stool samples from *C. jejuni*-positive chickens deposited in sun-exposed patios still had viable organisms demonstrable after 24 hours. At 48 hours, only 18 percent (10/57) of the samples were positive.

Discussion

Campylobacter jejuni infection most commonly occurs through human fecal-oral transmission or through contaminated food or water.[9–13] This transmission is difficult to document, unless direct observation of family hygienic practices is used. Stanton, *et al*, demonstrated that responses to 24-hour recall and knowledge-attitude-practices questionnaires do not reliably reflect observed hygienic household practices.[5] The direct observation method is a strength of this study. A limitation is that the sample size is small. These data nonetheless demonstrate that contamination of toddlers' hands with feces deposited in the house by free-roaming birds is extremely common in Peruvian pueblo jovenes. In Bangladesh, play areas of infants and toddlers were also seen to be contaminated with chicken feces in approximately 66 percent of the households.[14] Approximately one-half the Bangladeshi mothers claimed to have seen the child touching or eating animal feces during the previous two weeks.

Unwashed hands of toddlers in a pueblo joven are frequently contaminated with poultry feces. As 50 percent of the chickens in a pueblo joven are carriers of *C. jejuni*,[4] the health risk to these toddlers is evident. We found that few of these feces-to-hand episodes were interrupted by hand washing.

Many pueblo joven mothers do not believe that chicken stools provoke illness. Similar results were found in the Bangladesh study, where over one-half of the mothers also did not recognize animal feces as disease-producing.[14]

The absence of running water in pueblo jovenes and the high cost of the water used may be important factors discouraging hand washing. Even in developed countries where water is easily available, oral-fecal contamination is common. Direct observations in an English nursery school during an epidemic of shigellosis showed that 50 percent of the children had contaminated their hands with *Shigella sonnei* after using the school toilet. Moreover, one-third of the children handled their face or sucked their fingers after using the toilet.[15]

The risk of acquiring viable *C. jejuni* in the study area is augmented because the house floors are usually made of dirt or unfinished concrete, so the complete removal of poultry feces is difficult to achieve, and *C. jejuni* can survive in the ambient environment for several days. Even where water is plentiful, in homes with free-roaming chickens, it is not realistic to assume that mothers will be able to effectively interrupt transmission by handwashing.

In addition to *C. jejuni*, a small percentage (2 percent) of poultry were fecal carriers of *Aeromonas* species.[4] The relationship of this organism to diarrhea has not been established. None of the free-roaming chickens sampled excreted *Salmonella* species. In both Peru and Bangladesh,[1,16] *Sal-*

monella species is an infrequent cause of community epidemics of diarrhea.

The present study suggests similar socioeconomic background and characteristics of families with and without poultry in a pueblo joven population, making it difficult to identify a target population. In addition, an intervention program would have to overcome common beliefs that inherently encourage the transmission of *C. jejuni*. Many families raised uncorralled chickens even though they recognized that purchasing poultry was cheaper than home grown animals, believing that chickens grow better when they are corralled; and appreciated the risk in having chickens roaming freely through the house. This suggests that this behavior will be difficult to modify.

In Peru, as our data show, most individuals living in pueblo jovenes have access to mass media. Most children attend primary school. Both mass media and formal education in the school system could be utilized for intervention programs. These educational programs should emphasize the health risks to toddlers of raising free-roaming poultry within the household and the need to corral poultry and prevent their entrance into the house.

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8th Competition Announced for Health Professions Students

Dr. Louis W. Sullivan, Secretary of Health and Human Services, has announced the 8th annual national competition among health professions students for the "Secretary's Award for Innovations in Health Promotion and Disease Prevention." Winners will receive awards of $3,000 for first place, $2,000 for second, and $1,000 for third; up to 17 semifinalists will receive $250 each.

The competition is open to a broad range of health professions students enrolled in baccalaureate or higher degree programs affiliated with the Federation of Associations of Schools of the Health Professions, cosponsor of the competition. Those eligible to compete include students of medicine, osteopathic medicine, nursing, veterinary medicine, optometry, pharmacy, podiatric medicine, public health, health administration, health education, and allied health.

Papers should present innovative proposals for health promotion or disease prevention in the community, at the worksite, in an educational setting, or for a special population group.

The deadline for submitting entries is March 15, 1990. A paper must be submitted to the school attended by the author or authors. Papers will be judged first by the school and then by the professional associations. A departmental panel will make the final selections, and HHS Secretary Sullivan will announce the winners in July. For further information, contact Blake Crawford, Health Resources and Services Administration (HRSA), US Public Health Service, DHHS, 5600 Fishers Lane, Rockville, MD 20857; tel: (301) 443-3376.

HIV Seropositivity of Needles from Shooting Galleries in South Florida

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Abstract: Needle/syringe combinations were collected from three shooting galleries in South Florida and tested for the presence of antibodies to HIV-1. Fifteen of 148 needles (10.1 percent) tested positive for HIV-1 antibody. Seropositivity rates did not vary by the day of the week of collection, nor by shooting gallery from which they were collected. When the needle appeared to contain blood residue, 20.0 percent were positive versus 5.1 percent with no blood residue. These findings suggest that needles/syringes used in shooting galleries are likely to serve as reservoirs and/or vectors of transmission of the HIV-1 virus, and that although visual inspection of the needle/syringe may be useful in lessening the chance for transmission, even the visually "clean" needles may result in transmission of infection. (*Am J Public Health* 1990; 80:150–152.)

Introduction

The self-reported sharing of needles and syringes, particularly those belonging to shooting galleries, has been consistently associated with increased risk for HIV-1 infection among parenteral drug users.[1–3] A shooting gallery is a place where addicts go to rent syringes and needles ("works") and to inject drugs.

Shared drug injection equipment can transmit HIV-1 by parenteral transfusion when residual, contaminated blood remains in previously used syringes and needles and that equipment is reused by another person. Blood residue is often present in the syringe because of the aspiration of venous blood into the syringe.

No one has assessed the extent to which equipment owned by shooting galleries is positive for HIV-1 antibodies. Recently a study of needle-and-syringe exchange programs in Sydney, Australia determined that antibodies for HIV-1 were present in 3.1 percent of 1,544 needle and syringe combinations exchanged at two exchange centers.[4] Risk of exposure to HIV-1 probably is higher in shooting galleries where a wide variety of people rent the same works and needle and syringe combinations are routinely reconditioned to extend their useful life far beyond the average nine uses reported by intravenous users in Miami who reuse or share personal works.

In order to examine the potential for HIV-1 transmission through the use of injection equipment available in high-risk settings, needle/syringe combinations were collected from shooting galleries frequented by parenteral drug users in Miami, Florida and were tested for antibodies to HIV-1. Although this study measures antibody and not the presence of an infectious dose of virus, these data help explain the association between shooting galleries and HIV-1 infection.

Methods

Selection of Shooting Galleries

Three shooting galleries located in Miami, Florida were selected as the sites from which to collect needle/syringe combinations. These were among the most frequently mentioned galleries that were attended by parenteral drug users enrolled in coordinated investigations of the natural history of HIV-1 and the evaluation of risk behavior reduction programs. Seroprevalence* among users enrolled in these studies is approximately 36.1 percent for Blacks who were not Hispanic, 30.2 percent for Hispanics,** and 8.9% for Whites who were not Hispanic. The majority of persons who frequented these galleries were Blacks who were not Hispanic. Each gallery was located in a different inner-city area known for high rates of drug use.

Field observations at eight additional galleries indicate that the modes of operation at the study sites were typical for area shooting galleries. Each was housed in the residence of the gallery operator. All three rented used works for $2 and had bleach or alcohol available for cleaning works, and none sold drugs. Heroin and/or cocaine, both available for purchase nearby, were predominant. Syringe and needle repair and cleaning occurred in these galleries but no observations were made specifically for these needles and syringes.

Needle/Syringe Collection

Access to the three galleries was gained through the efforts of a staff outreach worker who had established a variety of contacts with the intravenous drug using population. He was a follow-up worker who spent all of his time locating drug using participants of a longitudinal study. Because he was not an outreach worker in an HIV prevention program, the shooting gallery operators were less likely to feel pressure to manipulate the selection of syringes. Needle/syringe units that had been used in the shooting galleries during the previous 24-hour period were collected each morning by the follow-up worker during two one-week periods in 1988. In order to avoid possible bias by day of the week, no more than six units were collected per site per day. Wherever more were available, six were selected by the follow-up worker without regard to appearance or condition.

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*Seroprevalence is based upon data from three coordinated studies of HIV infection among intravenous drug users (N = 1364) in Miami, Florida.

**The study population includes Hispanics who are White and Hispanics who are Black; they are Hispanic in culture and do not identify themselves as White or Black.

Each gallery operator was paid a flat fee of $24 per visit and did not participate in the selection of syringe units.

Each morning after collection, the needle and syringe combinations were immediately transported to the University of Miami and visually graded as follows:

1. "Visible Blood"—The needle/syringe appeared to contain liquid or dried blood.
2. "Dirty, No Visible Blood"—The needle/syringe contained visible dirt, or stains, but no blood.
3. "Clean, No Visible Blood"—The needle/syringe contained no visible dirt, stains or blood.

Needles were then delivered to the laboratory where they were tested for the presence of antibodies to HIV-1.

Laboratory Methods

A pilot study had been conducted with 17 needle and syringe combinations under conditions simulated to represent the condition of the needles/syringes which would be collected at the shooting galleries. Using the blood samples from a high-risk group whose blood was sent to the laboratory for routine testing, 200 μl of blood from each sample were drawn into a syringe and expelled. After standing overnight each needle/syringe was rinsed with 200 μl of saline and flushed 10 times with the saline solution. This solution was assayed twice by ELISA (Abbott Enzyme Immunoassay).

Ten samples from the test needle/syringe combinations were found to be reactive for HIV-1 antibodies confirmed by Western Blot and seven were non-reactive. These results were identical to the results of the routine testing for HIV-1 antibodies done on the blood samples.

Needle/syringe combinations collected from shooting galleries likewise were rinsed with 200 μl of saline, flushed 10 times and assayed in the same manner as the test needles. These procedures were similar to those used by Wolk[4] except that needles were tested more quickly and less saline was used in our study.

Needle/Syringes Not Tested

The field worker collected needle/syringe units regardless of condition or potential for testing or reuse. Sixty-two of the units collected could not be tested because of physical damage to the syringe (58 percent) or a clogged/damaged needle (48 percent). Needles and syringes were collected for three days during a subsequent week until 50 testable combinations had been received from each site. In two cases, results of HIV-1 testing were indeterminate on Western Blot, and were excluded from the analysis.

Analytical Procedure

Differences in proportions and 95 percent two-sided confidence intervals around the difference were calculated.

Results

Fifteen (10.1 percent) of the 148 needle/syringe combinations tested were found to be positive for HIV-1 (Site 1 = 10.2%; Site 2 = 12.3%, Site 3 = 8.0%) and 133 were negative (89.9 percent). There were no differences in seropositivity by day of the week on which needles were collected.

The relation between the graded condition of the needle/syringe combinations and the presence of HIV-1 antibodies is presented in Tables 1 and 2. Twenty percent of the equipment combinations which had visible remnants of blood had evidence of HIV-1 antibodies, compared to 5.1 percent in combinations without blood. Thus the risk of having an infected needle/syringe was approximately four times higher if the needle/syringe appeared to be bloody. No differences in seropositivity were observed when syringe and needle combinations which contained no visible blood were partitioned into clean (4.7 percent) and dirty (5.5 percent) categories (Table 2).

TABLE 1—Seropositivity of Needle/Syringe Combinations without and with Visible Blood

Appearance	Total Tested	Proportion HIV Ab+*
	N	%
No Visible Blood	98	5.1
Visible Blood	50	20.0
Total	148	10.1

*Difference in proportions 0.149 (0.06, 0.21)

TABLE 2—Seropositivity and Visual Appearance of Needle/Syringe Combinations without Visible Blood

Appearance	Total Tested	Proportion HIV Ab+*
	N	%
Clean	55	5.5
Dirty	43	4.7
Total	98	5.1

*Difference in proportions 0.008 (−0.06, 0.12)

Discussion

This study confirms epidemiological evidence that the use of needles belonging to shooting galleries can transmit HIV-1. Ten percent of the needle-syringe combinations which were tested from three shooting galleries were positive for antibodies to HIV-1. The use of such needles presents a clear public health hazard.

Shooting galleries provide needle and syringe combinations for "rent" to drug users. The combinations usually are held by the gallery operator or laid out on a table and are selected for use by either the operator or customer depending on house rules. Water and in some cases, alcohol or bleach, are available to rinse the works, i.e., to clean the syringe and needle. The custom in many galleries is to clean the works after injecting—not prior to injecting. However, there are no mechanisms to assure that works are cleaned after use and in many instances we found they were not cleaned. Moreover, cleaning often entailed simply rinsing with water rather than bleach or alcohol.

The data from this study suggest that the choice of clean-appearing injection equipment does not eliminate the possibility of using an HIV-1 infected needle. Needle and syringe combinations which contain visible blood were more likely to be positive for antibodies to HIV-1 than those that do not contain visible blood, but the selection of a needle without visible blood was no assurance of safety. This is important because drug users may mistakenly believe that they are capable of determining which used works are "clean" i.e., safe.

Works rented in most shooting galleries are returned for future use until they are no longer functional. In this study, 29.2 percent of the collected needle/syringe combinations were clogged and could not be tested. No attempt was made to test these non-functional needle/syringe combinations because of the risk of a needle stick or contamination by laboratory personnel (who would have had to remove needles to test the contents of the clogged needle or syringe). However, 60 percent of the untested combinations contained blood residue while only 33 percent of those tested contained visible blood. Thus, our estimate of seropositive needle-syringe combinations probably is a conservative estimate of seropositivity in usable needle/syringe combinations which at the time they were collected were available for use by shooting gallery customers.

Our choice of shooting galleries was not by random sample and does not purport to represent all shooting galleries in the area. For obvious reasons, it is not possible to draw up an inclusive list of shooting galleries from which to randomly select representative galleries. However, those studied were three of the largest and most frequented galleries in the South Florida area. The operators of each gallery reported they serviced an average of 125 intravenous drug users per week and many of these clients rented works on more than one occasion.

As with other drug use issues, the solution to the problem posed by HIV-1 contaminated needles and syringes is not simple. Intravenous drug users frequent shooting galleries for a variety of reasons.[5,6] In communities such as Miami where new syringes may be purchased by prescription only and possession of drug paraphernalia is illegal, galleries provide the user with a place to inject which is relatively safe from police intervention. Shooting galleries increase the availability of needle and syringe combinations and some also sell drugs. Galleries provide a place where drug users can socialize with other drug users. At some locations prostitution is common. Even faced with evidence that shooting galleries are sources of HIV-1 infection, many users still frequent galleries.

Intervention programs must be developed to reach intravenous users who continue to use shooting galleries. Both the users and gallery operators need to be educated on the necessity of either using new works or cleaning works and need to be taught proper cleaning techniques. Gallery operators exercise control over injection behavior at their locations. If they become convinced that sterile needles and syringes are essential, those central figures could have considerable influence upon a large number of other users. For example, two gallery operators in Miami have been observed requiring all customers to purchase new works.

If risk reduction is to occur among users who continue to inject drugs, techniques to clean works must be incorporated into the value system of the operators of the galleries as well as the individual parenteral drug users and made a part of the shooting behavior of those users who frequent shooting galleries.

ACKNOWLEDGMENTS

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NIH Consensus Statement on Oral Complications of Cancer Therapies

On April 17–19, 1989, the National Institute of Dental Research (NIDR) and the Office of Medical Applications of Research (OMAR) of the National Institutes of Health (NIH) held a consensus development conference on *Oral Complications of Cancer Therapies: Diagnosis, Prevention, and Treatment*. After hearing presentations on the subject by scientific experts, a consensus panel issued its report containing the conclusions and recommendations of the conferees. Free single copies of the consensus statement may be obtained from:

William H. Hall
Directior of Communications
OMAR, NIH, Bldg. 1, Room 259
Bethesda, MD 20892
Tel: (301) 496-1143

OMAR will also provide multiple copies of a two-page summary of the consensus statement to organizations that wish to distribute it to members of their organization. Contact Mr. Hall's office and advise him of the number of copies required.

Occupational and Other Exposures Associated with Male End-Stage Renal Disease: A Case/Control Study

N. Kyle Steenland, PhD, Michael J. Thun, MD, MPH, C. William Ferguson, BA, and Friedrich K. Port, MD

Abstract: We conducted a case-control study of 325 men ages 30–69 who were diagnosed with end-stage renal disease (ESRD) between 1976 and 1984, and resided in four urban areas of Michigan in 1984. Cases were selected from the Michigan Kidney Registry and excluded men with diabetic, congenital, and obstructive nephropathies. Controls were selected by random-digit dialing and were pair-matched to cases for age, race, and area of residence. Telephone interviews were conducted with 69 percent of eligible cases and 79 percent of eligible controls. Risk of ESRD was significantly related to phenacetin or acetaminophen consumption (odds ratio(OR) = 2.66), moonshine consumption (OR = 2.43), a family history of renal disease (OR = 9.30); and regular occupational exposures to solvents (OR = 1.51) or silica (OR = 1.67). Particular occupational exposures with elevated risk were solvents used as cleaning agents and degreasers (OR = 2.50) silica exposure in foundries or brick factories (OR = 1.92), and silica exposure during sandblasting (OR = 3.83). Little or no trend of increased risk with duration of exposure was found for these occupational exposures, with the exception of silica in sandblasting. Limitations of these data include representativeness of cases, possible overreporting by cases, and misclassification of exposures inherent in self-reports. (*Am J Public Health* 1990: 80:153–159.)

Introduction

End-stage renal disease (ESRD) is a major public health problem in the United States. In 1985, over 120,000 Americans suffered from ESRD, and each year approximately 30,000 new cases arise. Cases require dialysis or transplant, and the current annual cost exceeds $2 billion a year, almost entirely paid for by the federal government.[1] Yet very little is known about the causes of ESRD. A minority of ESRD cases are characterized etiologically (e.g., lead nephropathy, analgesic nephropathy, diabetic nephropathy).[2] Many cases are described histologically (e.g., glomerulonephritis, interstitial nephritis). Others are described as "hypertensive nephrosclerosis," with no determination of whether the kidney damage preceded or followed the hypertension.

Acute exposures to heavy metals can cause acute renal disease or dysfunction. There is also some indication that occupational exposures to metals, solvents, and silica may play a role in chronic renal disease.[2] Evidence indicates excess renal disease among workers exposed chronically to uranium,[3] cadmium,[4] and lead.[5–8] In addition, animal studies have shown that high exposures to carbon tetrachloride or perchloroethylene can also cause acute kidney damage.[9] In recent years, case-control studies have indicated that chronic glomerulonephritis is associated with occupational exposure to hydrocarbons.[10–15] Evidence of silica-related renal disease is limited to case reports[16,17] of acute renal failure following high exposures to silica.

To further investigate the hypothesis that occupational exposures might be associated with end-stage renal disease, we have conducted a case-control study of males diagnosed with ESRD between 1976–1984 in four urban areas of Michigan.

Methods

Cases

Patients were identified from the Michigan Kidney Registry,[18,19] which registers all new cases of ESRD in Michigan, collects demographic data, and records the diagnosis provided by the attending physician at the time of the ESRD diagnosis. Eligible cases were all men diagnosed between 1976 and 1984 with selected diagnoses of ESRD, ages 30–69 at diagnosis, and living in four urban areas in southeastern Michigan in 1984. Cases were restricted to those still living because of the unreliability of next-of-kin for reporting detailed exposure history. Diagnoses excluded were diabetic nephropathy, polycystic kidney disease, heroin nephropathy, lupus nephropathy, nephropathy due to malignancy, Alport's syndrome, unspecified chronic renal failure, nephropathy due to obstruction, and a variety of less common nephropathies grouped as "other." These diagnoses were excluded because most have known nonoccupational causes. The diagnoses included were glomerulonephritis, hypertensive kidney disease (nephrosclerosis), and interstitial kidney disease (including pyelonephritis, lead nephropathy and other unspecified interstitial disease). These diagnosis included represented about 65 percent of the ESRD patients in Michigan.

The summary diagnoses in the Michigan Kidney Registry may not be accurate, since many of ESRD patients present with advanced disease, and only an estimated 5 percent or less of ESRD patients have been biopsied. For example, some physicians may have classified a case as nephrosclerotic on the basis of high blood pressure at the time of diagnosis, without further data indicating that the high blood pressure preceded the kidney disease. While the diagnoses available to the Registry were sufficient to exclude certain well-defined entities such as diabetic nephropathy, the diagnoses of the cases which were included (nephrosclerosis, glomerulonephritis, and interstitial nephritis) were not clearly differentiated one from another, and were therefore analyzed together.

The Michigan Kidney Registry identified 612 eligible cases. Contact was initially made by mail, with phone calls to nonrespondents. Eighty-seven of these men had died prior to any attempt to contact them, and 14 died after consenting to be interviewed but prior to being interviewed; another 26 men were found upon interview to be ineligible because they had kidney disease subsequent to diabetes or prolonged heroin use; and 14 men could not participate either because they were in prison or were mentally incompetent. Of the remaining 471 men, we interviewed 325 (69 percent). More nonre-

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From the National Institute for Occupational Safety and Health (Steenland, Thun) and the Michigan Kidney Registry (Ferguson, Port). Address reprint requests to N. Kyle Steenland, PhD, NIOSH, R13, Robert A. Taft Laboratories, 4676 Columbia Parkway, Cincinnati, OH 45226–1998. This paper, submitted to the Journal February 6, 1989, was revised and accepted for publication June 26, 1989.

spondents than respondents were Black (67 percent vs 44 percent) and were living in inner Detroit (61 percent vs 40 percent).

Controls

Controls were chosen from the general population in the four urban areas of the study, via random-digit dialing. Controls were pair-matched to cases on age (within five years), race (White versus other), sex, and residential area based on telephone prefix, grouped in 15 geographical areas covering the entire study area. Men who had chronic kidney disease were excluded.

Random digit-dialing was conducted according to a modified Waksberg method[20] from February to October 1985. A potential control identified in calling the randomized number was allowed to match any case of the same age and race in the same geographical area. If numbers did not answer, repeated phone calls were made varying the time of day and day of week prior to abandoning the phone number.

A total of 3,962 numbers were called (with an average 1.95 calls per number). Of these, 61 percent (2,430) were working residential numbers. Twenty-three percent (567) of those who answered at a residential phone refused to provide initial information about males in the household. Once a household did provide such initial information and proved to have an eligible male, 79 percent (325/410) agreed to be interviewed. We compared the 325 controls (respondents) with the 85 individuals who were eligible but refused the interview (refusals). Potential controls who refused differed from respondents regarding race (55 percent Black versus 44 percent), and residence (60 percent from inner Detroit, versus 40 percent).

Interviews

Telephone interviews lasting 30–45 elicited demographic, medical history, and occupational histories. Cases were interviewed from November 1983 through December 1985, while controls were interviewed from March 1985 to January 1987. Interviewers knew which respondents were cases from the answers to a variety of questions. Neither cases nor controls were aware of the hypotheses of the study.

Respondents were asked about regular moonshine use, regular pain pill use, family history of serious kidney disease, years of education, smoking, lead poisoning, injected antibiotics, and environmental exposure to metals and solvents. Exposures were truncated at the year of diagnosis as end-stage for cases and their matched controls. Phenacetin or acetaminophen use was determined by a review of answers to a question about the use of pain pills on a regular basis (more than one pill per week for two years or more). Brands of analgesics were classified by whether or not they contained phenacetin or acetaminophen in the early 1970s, based on data provided by Dr. Dale Sandler at the National Institute of Environmental Health (personal communication). Acetaminophen was grouped with phenacetin because it is a metabolite of phenacetin and may also be associated with renal disease.[21]

Family history of renal disease was considered positive if kidney disease (excluding stones and cancer) occurred among parents, children, or siblings, and was not known to have occurred subsequent to diabetes or trauma.

Each job held for more than six months since age 18 was classified for occupation and industry using 1980 US Census codes.[22] For each job, questions were asked regarding regular exposure to solvents, metal particles, metal fumes, mercury, oil and gasoline, and ammonia. Ammonia was included as a check, since it was not thought to be related to kidney disease. For each material, details of the exposure (e.g., hours per week, process description) were recorded. Work history for cases and their matched controls was considered only up until the year of the diagnosis as end-stage for the case.

We reviewed the completed interviews without knowledge of case-control status and with the assistance of NIOSH (National Institute for Occupational Safety and Health) industrial hygienists. For positive responses to questions about metal fumes, metal particles, solvents, and silica, we created sub-categories by type of solvent, metal, etc (Table 2). When an individual had been unable to recall the number of hours per week during which he was exposed, we arbitrarily assigned five hours a week for infrequent exposure, 20 hours a week for frequent but not constant exposure, and 40 hours a week for constant exposure. In a few cases, a self-reported exposure to a substance was clearly wrong due to misclassification by the respondent, and we corrected the response (e.g., self-reported exposure to solvents such as "lime" or "pesticides"). In other cases, we called a respondent back for clarification.

Analysis

Odds ratios (OR) were estimated from conditional logistic regression.[23] Odds ratios for exposure were adjusted for confounding by nonoccupational variables and other occupational variables. Dummy variables (missing versus not exposed) were added to the model to account for those instances in which study subjects were missing data for lifestyle variables. When subjects were missing data on exposure, they were called again to resolve the problem. No cases or controls were lost at the analysis stage due to missing data.

We compared the occupational history of cases and controls in several ways:

- We considered a history of having ever been exposed to specific substances (e.g., all solvents, or solvents used as degreasers).
- We estimated cumulative exposure by multiplying the number of hours per week a man was exposed by the number of years in the job for each job, and summing over all jobs.
- Finally, for solvents, we used an exposure matrix developed by Shalat and others[24] to identify jobs likely to have involved high exposure to solvent (see Appendix). Cases and controls were then compared regarding work in these jobs. Analyses were also run which discounted any exposures (lifestyle or work history) which occurred in the five years preceding the date of diagnosis of the case.

Results

Table 1 gives the frequency of a number of lifestyle variables among (unmatched) cases and controls. After controlling for confounding by years of education and date of birth, phenacetin/acetaminophen use and moonshine use were both significant predictors of kidney disease, with odds ratios of 2.66, (95% confidence intervals 1.04, 6.82) and OR 2.42 (95% CI 1.10, 5.36), respectively. Cumulative measures of exposure for these variables showed positive but nonsignificant trends of increased risk with increased duration of use. The findings for family history were the most remarkable, with cases much more likely than controls (OR 9.30, 95% CI 7.99, 10.82) to have a close family member (parent,

TABLE 1—Demographic Characteristics of Study Population

Characteristics	Cases (n = 325)	Controls (n = 325)
Average years of education	11.7	12.5
Number non-White*	143	143
Number living in inner Detroit	130	130
Number living in outer Detroit	114	114
Number living in Flint, Saginaw, or Lansing	81	81
Some college	106	129
Average date of birth	1930.0	1930.5
Family kidney disease	37	7
Regular moonshine use	31	10
Regular use of phenacetin/acetaminophen	22	7
Current cigarette smokers**	143	151
Former cigarette smokers	119	91
Average years smoked, current smokers	35.8	34.4

*93% of non-Whites were Black.
**Includes men who quit within one year prior to their diagnosis.

child, sibling) with kidney disease (excluding kidney stones and kidney disease subsequent to trauma or diabetes). When any exposure within five years of the date of diagnosis of the case and his matched control was ignored, the results for moonshine consumption were unchanged, while the odds ratio for phenacetin/acetaminophen fell slightly from 2.66 to 2.47 (0.86, 7.12).

Controls were more likely ever to have worked in jobs as professionals, managers, or technicians (56 percent vs 46 percent). The most common industry for both cases and controls was the automobile industry, in which 51 percent of cases and controls had worked. The next most common was construction (17 percent), followed by foundries (9 percent), and hospitals (6 percent).

Table 2 presents the results of a logistic regression for occupational variables. Adjusted odds ratios were significantly elevated for exposure to solvents used as degreasers and silica used in foundries or in sandblasting. An elevated odds ratio was also observed for exposure to lead. No interaction terms were significant between any exposure variables and any lifestyle variables or other exposure variables.

Most men exposed to solvents used as degreasers worked in metal manufacturing industries (Census codes 271–370, 47 cases, 22 controls). Within this sector, the principal industry was auto manufacturing (32 cases, 17 controls). About equal numbers of cases and controls were exposed in dry cleaning shops (six cases, five controls). The subcategories of silica with elevated risk were exposure in foundries, brick making, or sandblasting, which were the categories likely to have involved the highest exposures.

There were no strongly positive trends of increased risk with increased duration of exposure for any exposure variable, with the exception of silica used in sandblasting. The odds ratio for exposure to silica in sandblasting for 2.5 years for 40 hours a week (cumulative duration = 100, an average exposure) was 2.74 (.69, 9.97).

We also considered employment in jobs considered to have had high solvent exposure (see Appendix). There were no differences between cases and controls regarding either ever having worked in these jobs (31 cases, 30 controls) or the length of time having worked in these jobs (5.3 years for cases, 6.8 for controls).

The above results for occupational variables were changed little when all exposures in the five years preceding the date of diagnosis of the case (and matched control) were ignored. For example, the odds ratio for ever having been exposed to solvents increased from 1.51 to 1.68 (1.13, 2.48), while the odds ratio for ever having been exposed to silica increased from 1.67 to 1.81 (1.09, 3.02). The coefficients for cumulative duration of exposure also were affected only slightly (for solvents, the coefficient changed from −.0000095 to .00009, while for silica it changed from −.0039 to −.0004).

Discussion

This study suffered from several limitations inherent in an interview study of ESRD patients. First, patients are often

TABLE 2—Adjusted Logistic Regression Results for Occupational Variables

Exposure	Number Cases Exposed	Number Exposed	Odds Ratio* (95% CI) Ever-Never Exposed*
All solvents	124	82	1.51 (1.03, 2.22)
Solvents used in paints and glues	38	34	1.01 (.58, 1.74)
Solvents used as cleaning agents or degreasers	94	40	2.50 (1.56, 3.95)
Solvents used in other processes	17	13	1.05 (.44, 2.48)
Metal fumes	139	94	1.17 (.77, 1.80)
Lead (soldering and other)	32	16	1.73 (.82, 3.65)
Iron/steel	109	72	1.15 (.73, 1.80)
Welding fumes	51	44	0.75 (.44, 1.28)
Metal particles	119	96	0.97 (.58, 1.48)
Oil and gas	135	129	0.74 (.64, 1.51)
Gas and diesel fuel	74	65	0.98 (.49, 1.06)
Motor and fuel oil	68	49	1.13 (.69, 1.84)
Cooling fluids	26	40	0.50 (.27, .92)
Silica	87	54	1.67 (1.02, 2.74)
Cement and sand	17	18	0.78 (.34, 1.78)
Brick and foundry	57	31	1.92 (1.06, 3.46)
Other silica	16	12	1.08 (.42, 2.77)
Sandblasting	9	3	3.83 (.97, 15.19)
Ammonia	33	19	1.31 (.66, 2.60)

*Models included variables for date of birth, years of education, regular use of moonshine, and family history of kidney disease, metal fumes, solvents, and silica.

quite ill and difficult to interview. This may contribute to a high refusal rate. Second, 482 men had been diagnosed between 1976–84 and would have been eligible for our study but had died by 1984, so that our study was of a population of survivors. A demographic analysis of those who died before they could enter the study indicated they differed little on race or residence, but were approximately five years older than men who were studied. A bias in our results could exist if these decedents differed substantially in their occupational exposures with those who survived to be studied. It is possible, for example, that those who died were of lower socioeconomic status and had different exposures than those who survived, although it is not clear that this would either increase or decrease odds ratios. We did not attempt to use of surrogate interviews by next-of-kin because next-of-kin could not report the decedent's detailed occupational histories. Third, it is difficult to study separate diagnostic groups among ESRD patients. We did not have confidence that the summary diagnosis in the Registry was sufficiently accurate to allow us to divide our data into the principal diagnostic groups (glomerulonephritis, nephrosclerosis, pyelonephritis) for diagnosis-specific analyses. Many of the above difficulties could be overcome by studying individuals with less severe kidney disease, prior to end-stage. With little data available to date, it is not clear whether occupational associations are more or less apparent in patients with end-stage versus less severe disease.

As in many case-control studies, we relied on self-reports to determine exposures. Cases may have had a tendency to overreport exposures compared to controls. We included a variable (ammonia exposure) thought not to be related to renal disease to test such overreporting. Cases did report more exposure to ammonia than did controls, although numbers were small and the difference was not statistically significant. Complicating any assessment of overreporting among cases was the fact that cases were more likely to have had less education than controls, and were more likely to have worked in blue-collar jobs than controls. Hence, they were more likely to have truly been exposed to occupational agents. This increased likelihood of "blue collar" status among cases compared to controls may have been an artifact of the random-dial dialing process. An alternative explanation is that our findings may represent a real phenomenon, whereby people with renal disease are of a lower socioeconomic strata than those without renal disease. While the published literature provides little data to support this thesis, one recent report reached similar conclusions.[21] To control for the observed confounding effects of socioeconomic class in our study, we have included a variable in the model for years of education.

Regarding nonoccupational variables, the effect of moonshine on renal disease is presumably due to contamination by lead.[25] Other studies have indicated that analgesics, particularly phenacetin, may be associated with renal disease (see the consensus report[26]).

Regarding occupational variables, our findings of elevated risk associated with self-reported exposure to solvents (used as cleaning agents or degreasers) and silica (used in foundries or brick manufacturing, and in sandblasting) are consistent with reports of renal disease resulting from acute exposure to these substances. An effect of solvents on chronic renal disease has also been seen in case-control studies.[15]

The lack of trend with duration of exposure to solvents can be seen as weakening the observed association between solvents and ESRD. However, duration of exposure may be a poor surrogate for cumulative dose, if those with short-term exposures also have higher exposures. Similarly, the negative findings for solvents using the pre-defined "high exposure" jobs may be due to the insensitivity of the job-exposure matrix which we used. Few attempts have been made to validate job-exposure matrices (none had been attempted for the one we used), and when such attempts have been made results have not always been encouraging.[27]

The covariate with the strongest association with renal disease in our study was a history among close relatives (parents, children, siblings) of renal disease (other than stones, or kidney disease subsequent to cancer or trauma). Due to the extremely strong association (odds ratio 9.3) observed for this variable, it seems unlikely that this positive association was due to overreporting. Familial patterns are known to exist for a number of relatively rare kidney diseases, but many of these would have been excluded from our case series (e.g., Alport's syndrome, polycystic kidney disease). Family history of hypertension, however, was not examined in our study, and may have contributed to the positive association we observed.

In summary, this study found a number of associations with ESRD which, in our opinion, warrant further investigation. These associations were generally supported by other data in the literature. The epidemiology of chronic renal disease is a relatively new area of investigation with many potential benefits, in light of the high prevalence of renal disease, its high mortality rate, and its high cost of treatment.

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APPENDIX

*US Census Codes for Jobs Considered Likely to Have Had High Exposures to Solvents**

Occupation Code/Occupation	Industry Code	Industry
203 Clinical lab tech	any	any
224 Chemical tech	any	any
455 Pest control	any	any
188 Painter, sculptor	any	any
789 Handpainting, coating, decorating	any	any
579 Painters, constr and maint	any	any
736 Typesetters	any	any
734 Printing machine operators	any	any
759 Painting/paint spray machine opr	any	any
885 Garage servs/gas station attnd	any	any
764 Washing, cleaning, pickling	270–301	Metal
725 Misc. metal/plastic processing machine operator	270–301	Metal
756 Mixing/blending machine opr	181,191,192,190	Chemicals, drugs, paints
	180,210,211,212	Rubbers and plastics
757 Separating filtering mach opr	181,191,192,190,	Chemicals, drugs, paints
	200,201	Fuel
758 Compressing/Compacting mach opr	180,210,211,212	Rubber, plastics
709 Grinding, abrasive, buffing, polishing	180,210,211,212	Rubber, plastics
705 Milling/planning mach opr	180,210,211,212	Rubber, plastics
777 Misc. machine opr	180,210,211,212	Rubber, plastics
755 Shaping/forming mach opr	180,210,211,212	Rubber, plastics
753 Cementing/gluing mach opr	180,210,211,212	Rubber, plastics
786 Handcuffing/trimming	180,210,211,212	Rubber, plastics
787 Hand molding/casking/forming	180,210,211,212	Rubber, plastics
796 Production inspectors	180,210,211,212	Rubber, plastics
889 Laborers	220,221,222	Leather
745 Shoe machine opr	220,221,222	Leather

*SOURCE: Shalat 1986. Shalat ranked all possible jobs on a scale of 0–3 for both likelihood and intensity of exposure. Jobs here defined as high solvent jobs were those for whom the product of intensity and exposure was five or more.

Sentinel Health Events (Occupational): Analysis of Death Certificates among Residents of Nassau County, NY between 1980–82 for Occupationally Related Causes of Death

JUDITH P. FELDMAN, MD, MPH, AND LINDA M. GERBER, PHD

Abstract: Death certificates for residents of Nassau County, New York dying between 1980 and 1982 were examined for causes of death defined as Sentinel Health Events (Occupational) {SHE(O)}. Of 16,193 deaths from all causes, 2,286 (14.1 percent) were identified as SHE(O) deaths; 142 required no match because the cause of death was inherently related to the occupation or industry. Malignant neoplasms of the trachea, bronchus, or lung were the most frequently occurring SHE(O), accounting for 60 percent of all SHE(O) deaths and 81 percent of matched SHE(O) deaths. The construction industry was associated with the vast majority of such deaths. (*Am J Public Health* 1990; 80:158–161.)

Introduction

Although accurate numbers of job-related injuries, illnesses, and deaths are difficult to obtain, estimates suggest that two to 10 million injuries, 14,000 accidental deaths, and 100,000 occupationally related illnesses result annually from workplace accidents and exposure.[1,2]

Recognition and surveillance of occupationally related deaths are difficult since only 28 of 50 states routinely code industry and/or occupation (I/O) data on death certificates,[2] and methods of record keeping, sorting and coding are not standardized.

Rutstein, *et al*, have suggested a system to identify preventable occupationally related diseases, or Sentinel Health Events (Occupational), [SHE(O)].[3] Recognition of SHE(O)s may aid in the control of occupational diseases by providing the impetus for epidemiologic or industrial hygiene studies, or serving as a warning signal that material substitution, engineering control, or personal protection of medical care may be required.

New York State is one of several states that routinely codes I/O data on death certificates. The New York State Department of Health had recently implemented coding occupational and industry entries on death certificates according to the US Bureau of the Census coding system,[4] and, at the time this study was undertaken, had completed coding on death certificates between 1980 and 1982.[5]

Nassau County on Long Island has a population of 1.3 million,[6] and a diverse industrial and occupational makeup.[7,8] Because of its large population and varied economic environment, a study was developed to implement Rutstein's SHE(O) system to analyze occupationally related disease patterns and examine whether any particular examples of occupations or industries put workers in Nassau County at increased risk for disease.

Methods

All death certificates recorded by the New York State Department of Health for residents of Nassau County whose age at death was between 18 and 74 years, and whose date of death was between January 1, 1980 and December 31, 1982, were examined. Death certificates had been assigned an ICD(9) code based on the diagnosis most likely to have been the underlying cause of death. In addition, a code was assigned for both the occupation and industry listed on the certificate.

Using Rutstein's list of occupations and industries, death certificates whose cause was on the SHE(O) list were sorted into five categories: 0 = unmatched to occupation and industry; 1 = match on industry only; 2 = match on occupation only; 3 = match on both occupation and industry; and 4 = no match needed since listed cause of death was considered to be inherently related to occupation regardless of listed occupation or industry (e.g., mesothelioma).

Categories 1–4 were grouped as "SHE(O) deaths with I/O match" and compared by sex, age, race, occupation, and industry to the total deaths from all causes and to the total SHE(O) deaths (matched and unmatched).

Results

The 50 SHE(O) categories suggested by Rutstein are listed in the Appendix. Only 26 were identified as the underlying cause of death on one or more death certificates in Nassau County (shown with an asterisk in the Appendix.) There were 16,193 deaths from all causes of which 2,286 (14.1 percent) were SHE(O) deaths and 142 (0.9 percent) were matched SHE(O) deaths.

Despite the fact that males made up only 47.8 percent of the population between ages 18–74, they accounted for 60.3 percent of deaths from all causes, 66.3 percent of all SHE(O) deaths, and 90.1 percent of matched SHE(O) deaths.

Malignant neoplasms of the trachea, bronchus, or lung (respiratory tract neoplasms) were the most frequently occurring SHE(O), accounting for 59.5 percent of all SHE(O) deaths. The other frequently encountered SHE(O)s were cirrhosis of the liver without mention of alcohol [10.4 percent of all SHE(O)s], malignant neoplasms of the bladder [5.2 percent of all SHE(O)s], and malignant neoplasms of the kidney [5.7 percent of all SHE(O)s] (Table 1).

Of the 2,286 SHE(O) deaths, 142 (6.2 percent) were matched; of those, 117 (82.4 percent) matched on industry alone, 13 (9.2 percent) required no match, 11 (7.7 percent) matched on occupation alone, and one (0.7 percent) matched on the combination of industry and occupation.

Although respiratory tract neoplasms accounted for the largest proportion of the SHE(O) deaths, only 8.4 percent of

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TABLE 1—Number and Percent of SHE(O) Deaths, and SHE(O) Deaths as Percent of Deaths from All Causes According to Cause and Sex: 1980–82

Underlying Cause of Death	# SHE(O) Deaths			% SHE(O) Deaths			% Deaths from All Causes		
	M	F	Total	M	F	Total	M	F	Total
Pulmonary TB	6	3	9	0.4	0.4	0.4	0.1	0.1	0.1
Chronic hepatitis	12	4	16	0.8	0.5	0.7	0.1	0.1	0.1
Cirrhosis liver without alcohol	142	97	239	9.4	12.6	10.4	1.5	1.5	1.5
Hepatitis-non-A, non-B	1	6	7	0.1	0.8	0.3	0.0	0.1	0.0
Hemangiosarcoma liver and site unspecified	39	44	83	2.6	5.7	3.6	0.4	0.7	0.5
MN nasal cavity	3	2	5	0.2	0.3	0.2	0.0	0.0	0.0
MN larynx	46	10	56	3.0	1.3	2.4	0.5	0.2	0.3
MN trachea, bronchus, lung	936	425	1361	61.8	55.1	59.5	9.6	6.6	8.4
Mesothelioma (all)	10	2	12	0.7	0.3	0.5	0.1	0.0	0.1
MN bone	9	2	11	0.6	0.3	0.5	0.1	0.0	0.1
MN bladder	94	24	118	6.2	3.1	5.2	1.0	0.4	0.7
MN kidney	88	42	130	5.8	5.4	5.7	0.9	0.7	0.8
Lymphoid leukemia	26	6	32	1.7	0.8	1.4	0.3	0.1	0.2
Myeloid leukemia	42	39	81	2.8	5.1	3.5	0.4	0.6	0.5
Erythroleukemia	2	1	3	0.1	0.1	0.1	0.0	0.0	0.0
Aplastic anemia	1	3	4	0.1	0.4	0.2	0.0	0.1	0.0
Agranulocytosis or neutropenia	3	0	3	0.2	0.0	0.1	0.0	0.0	0.0
Toxic effects of lead or its compounds	5	0	5	0.3	0.0	0.2	0.1	0.0	0.0
Extrinsic asthma	8	14	22	0.5	1.8	1.0	0.1	0.2	0.1
Coalworkers' pneumoconiosis	1	0	1	0.1	0.0	0.0	0.0	0.0	0.0
Toxic hepatitis	4	3	7	0.3	0.4	0.3	0.0	0.1	0.0
Acute renal failure	5	9	14	0.3	1.2	0.6	0.1	0.1	0.1
Chronic renal failure	11	17	28	0.7	2.2	1.2	0.1	0.3	0.2
Renal failure-unspecified	21	18	39	1.4	2.3	1.7	0.2	0.3	0.2
Total*	1515	771	2286	100.1	100.1	99.7	15.6	12.2	14.0

*percentages may not total 100.0 due to rounding

TABLE 2—Number and Percent of SHE(O) Matched Deaths by Underlying Cause

Underlying Cause of Death [SHE(O)'s] only	Total SHE(O) Deaths	SHE(O) Matched # Deaths (%)
Pulmonary TB	9	1 (11.1)
Chronic hepatitis	16	0 (0.0)
Cirrhosis of liver without alcohol	239	8 (3.3)
Hepatitis- non-A, non-B	7	1 (14.3)
Hemangiosarcoma of liver and site unspecified	83	0 (0.0)
MN of nasal cavity	5	0 (0.0)
MN of larynx	56	2 (3.6)
MN of trachea, bronchus or lung	1361	115 (8.4)
Mesothelioma, (all)	12	12(100.0)
MN of bone	11	0 (0.0)
MN of bladder	118	0 (0.0)
MN of kidney and urinary organs	130	0 (0.0)
Lymphoid leukemia	32	0 (0.0)
Myeloid leukemia	81	0 (0.0)
Erythroleukemia	3	0 (0.0)
Aplastic anemia	4	0 (0.0)
Agranulocytosis or neutropenia	3	0 (0.0)
Toxic effects of lead or its compounds	5	0 (0.0)
Extrinsic asthma	22	0 (0.0)
Coalworker's pneumoconiosis	1	1(100.0)
Toxic hepatitis	7	0 (0.0)
Acute renal failure	14	0 (0.0)
Chronic renal failure	28	1 (3.6)
Renal failure-unspecified	39	1 (2.6)
Total	2286	142 (6.2)

them were matched. Although 14.3 percent of non-A, non-B hepatitis, and 11.1 percent of tuberculosis cases were matched, the absolute numbers (one each) that were matched were very small. The 12 mesothelioma and one coal workers' pneumoconiosis were matched in 100 percent of the cases, but these diseases are by definition matched on diagnosis alone and fall into category 4 of "no match needed" (Table 2).

The 114 industry-matched respiratory tract neoplasms accounted for 80.3 percent of all matched deaths and 97.4 percent of all industry-matched deaths. Of the 13 deaths that required no match, 12 (92.3 percent) were due to mesothelioma (Table 3). For all respiratory tract neoplasms, the most commonly listed occupations were "housewives" (16.1 percent) and "managers and administrators not elsewhere classified" (8.4 percent); the most commonly listed industries were "private homes" (housewives) (16.2 percent) and "construction" (7.6 percent). For matched respiratory tract neoplasms, "construction laborers, plumbers, pipefitters and steamfitters, carpenters, construction supervisors not elsewhere classified" (7.8 percent) and "managers and administrators not elsewhere classified" (7.8 percent) were the most frequently reported occupations. "Construction" (87.0 percent), and "miscellaneous fabricated metal products" (5.2 percent) were the most frequently reported industries.*

Discussion

While the categories of SHE(O)s that were identified probably reflect the occupational, industrial, and environmental make-up of Nassau County, they might also reflect a lack of time or interest on the part of physicians in obtaining occupational histories. For example, despite the fact that there were 12 reported deaths from mesothelioma (virtually always caused by asbestos and usually easily recognized), there were no reported deaths from asbestosis (also caused by asbestos but not easily recognized in the absence of a good occupational history).

That men accounted for 60.3 percent and women only 39.7 percent of deaths from all causes can be explained by the fact that only ages 18–74 were analyzed. Women have a life expectancy at birth of 7.5 years greater than men.[9] Similarly, the fact that Whites accounted for 92.0 percent of all deaths

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*Additional details available on request to author.

TABLE 3—Underlying Cause of Death by Type of Match, SHE(O)s Only

Underlying Cause of Death	Total with Match	Industry Only	Occupation Only	Both I and O	No Match Needed
Pulmonary TB	1	0	1	0	0
Cirrhosis of liver without alcohol	8	1	7	0	0
Hepatitis-non-A, non-B	1	0	0	1	0
MN larynx	2	2	0	0	0
MN trachea, bronchus, or lung	115	114	1	0	0
Mesothelioma (all)	12	0	0	0	12
Coalworkers' pneumoconiosis	1	0	0	0	1
Chronic renal failure	1	0	1	0	0
Renal failure-unspecified	1	0	1	0	0
Total	142	117	11	1	13

can be explained by the fact that Whites make up 91 percent of the total population of Nassau County.[8]

Women accounted for a disproportionately small number of deaths in the SHE(O) category and in the matched SHE(O) category. Many of the diseases listed as SHE(O)s have a latency period of 10–30 years. Deaths in 1980 from SHE(O)s might well be a result of exposure in the 1950s and 1960s when only 25–30 percent of the women were in the labor force.[10] Moreover, many women are classified as "housewife" on their death certificates. Those who may have died from an occupationally related illness would not be identified through this system.

The most striking findings of this study are that while respiratory tract neoplasms accounted for only 8.4 percent of all deaths, they accounted for 59.5 percent of all SHE(O)s and 81.0 percent of the matched SHE(O)s, and that while the construction industry was associated with only 5.9 percent of all deaths and 6.6 percent of all SHE(O) deaths, it was associated with 72.1 percent of all matched SHE(O) deaths and 87.0 percent of all matched deaths from respiratory tract neoplasms. It is difficult, however, to interpret these findings. Although in the construction industry asbestos is a major risk factor for lung cancer, smoking is the single most important cause of these deaths, and enhances the carcinogenic risk of asbestos.[11] Since laborers tend to be heavy smokers,[12] the apparent high proportion of respiratory tract cancers among construction workers may be due to the increased rate of smoking in this population, as well as to asbestos exposure. Smoking history was not available on the death certificates.

Death certificates have other limitations. The accuracy of occupational data on death certificates when compared to information obtained from medical records, union records, and personal interviews has ranged from 51 to 93 percent.[13–16] Occupational data requested on the death certificate is for "usual occupation' " with space for only one entry. However, few people work in a single job or are exposed to a single agent in a working lifetime. Because of the latency period for many diseases, an early job with toxic exposure may not be the "usual job" listed. Occupational categories are general, and there is no information on actual job description, or working conditions.

Finally, because of the latency period for many diseases on the SHE(O) list, the use of the death certificates to evaluate occupational exposure gives information in many instances on exposure that occurred years ago. Since people move, exposure and death may not necessarily have occurred in the same geographic area, thus limiting the usefulness of death certificates in pinpointing occupational hazards except for substances causing diseases with relatively short latency periods.

As the number of potentially toxic substances added to the workplace grows, it becomes increasingly important to identify occupations and industries that pose a threat to workers' health. Rutstein's scheme is an innovative step toward that goal. However, the limitations of death certificates underscore the need for further work in this area, especially for improved reporting systems and data collection methods.

One promising step in that direction is the implementation by the New York State Department of Health of occupational registries for individuals in certain high-risk jobs, with exposure to known carcinogens, or with traces of certain heavy metals in their blood. All businesses employing 100 or more people are required to report annually to the State the names of those employees who fall into any of the aforementioned categories. Such individuals could then be followed for early disease that might be related to their on-the-job exposures.

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APPENDIX
ICD(9) Codes Searched as Sentinel Health Events (Occupational)+

Underlying Cause of Death	ICD(9) Code
* Pulmonary TB	0.11.0–011.9
Silicotuberculosis	502
Plague	020.0–020.9
Tularemia	012.0–012.9
Anthrax	022.0–022.9
Brucellosis	023.0–023.9
Tetanus	037
Rubella	056.0–056.9
Hepatitis A	070.0–070.1
Hepatitis B	070.2–070.3
* Chronic hepatitis	571.4
* Cirrhosis of the liver (non-alcoholic)	571.5, 571.8, 571.9
* Hepatitis- non-A, non-B	070.4–070.9, 573.5
Rabies	071
Ornithosis	073.0–073.9
* Hemangiosarcoma of the liver, and, site unspecified	155, 171.5, 171.9
* MM** of the nasal cavity	160.0, 160.1, 160.3, 160.4, 160.5, 160.8, 160.9
* MN of the larynx	161.0–161.9
* MN of the trachea, bronchus or lung	162.0–162.9
* Mesothelimoa	
* Retroperitoneum and peritoneum	158.0–158.9
* Pleura	163.0–163.9
* MN of the bone	170.0–170.9
MN of the scrotum	187.7–187.9
* MN of the bladder	188.0–188.9
* MN of the kidney and urinary organs	189.0–189.9
* Leukemia	
* Lymphoid	204.0–204.9
* Myeloid	205.0–205.9
* Erythroleukemia	207.0–207.8
* Aplastic anemia	284.8–284.9
Acquired hemolytic anemia	283.1–283.9
* Agranulocytosis or neutropenia	288.0, 288.9
Methemoglobinemia	289.7
Toxic encephalitis	323.7
* Toxic effect of lead and its compounds (including fumes)	984
Parkinson's disease-secondary	332.1
Cerebellar ataxia	334.3
Inflammatory and toxic neuropathy	357.7
Extrinsic allergic alveolitis	495.0–495.9
* Extrinsic asthma	493.0, 493.9
Detergent asthma	507.8
* Coalworkers' pneumoconiosis	500, 505
Asbestosis	501
Silicosis and talcosis	502
Chronic beryllium disease	503
Byssinosis	504
Acute bronchitis, pneumonitis and pulmonary edema due to fumes and vapors	506.0–506.9, 507.1
* Toxic hepatitis	570, 573.3
* Acute renal failure	584.0–584.9
* Chronic renal failure	585
* Renal failure, unspecified	586

+adapted from Rutstein *et al.*[4]

*causes of death and accompanying ICD codes that were found on at least one death certificate in Nassau County between 1980–82; SHE(O)s without an asterisk were not found on any death certificate.

**MN signifies malignant neoplasm.

A Statewide Case Registry for Surveillance of Occupational Heavy Metals Absorption

Michael E. Baser, MS, and David Marion, BA

Abstract: The New York State Heavy Metals Registry is a legislatively mandated program through which clinical laboratories, physicians, and health facilities report state residents 18 years of age and older with elevated levels of lead, mercury, arsenic, or cadmium in blood or urine. From 1982–86, the current employer was determined for 95.9 percent of 3,309 cases. Occupational exposures in 328 companies accounted for 82.8 percent of cases. The majority of companies were reported for lead (247 companies, 75.3 percent of total) or mercury (47 companies, 14.3 percent of total). Of the 247 companies reported to the Registry for lead, the Occupational Safety and Health Administration (OSHA) inspected 18 of 98 companies (18.4 percent) in the manufacturing sector, but only one of 149 companies (0.6 percent) outside the manufacturing sector. We conclude that the Registry effectively detects companies with heavy metals exposures, and is an especially useful adjunct to OSHA inspections outside the manufacturing sector. (*Am J Public Health* 1990; 80:162–164.)

Introduction

Heavy metals are common contaminants of the occupational environment. Absorption of metals can cause abnormalities of biochemical, cellular, and physiological function, as well as disease or death.[1] Occupationally related diseases are under-reported,[2] and surveillance may detect industries and companies with excessive exposures.[3–5] Methods of surveillance for occupational heavy metals exposure include Occupational Safety and Health Administration (OSHA) inspections for violations of the federal lead standard,[5,6] workers' compensation claims,[7] and the National Occupational Hazard Survey.[8,9]

In New York State, the two most useful sources of biological monitoring or industrial hygiene data are: 1) the New York State Heavy Metals Registry, a case registry of persons with elevated heavy metals levels in blood or urine, and 2) results of OSHA inspections for violations of the federal lead standard. The purposes of this paper are to present the methods and results of the Heavy Metals Registry during its first five years of operation (1982–86), and to determine whether companies reported to the Registry for lead have been inspected by OSHA for violations of the federal lead standard.

Methods

Sections 22.6 and 22.7 of the New York State Sanitary Code require that clinical laboratories, hospitals, and physicians report persons with heavy metals absorption to the state health department. The health department provides reporting forms which include name and residence, age, sex, heavy metals concentrations, report source (i.e., analytic laboratory, company, hospital, or physician), and private physician. Confidentiality is protected by a public health law which restricts use of the data to scientific or medical research and renders the information inadmissible as legal evidence.

Cases are defined from the reporting form as state residents 18 years of age or older with elevated levels of lead in blood (Pb-B $\geq$ 40 μg/dL or 1.93 μmol/L) or urine (Pb-U $\geq$ 80 μg/L), mercury in blood (Hg-B $\geq$ 5 μg/L or 0.02 μmol/L) or urine (Hg-U $\geq$ 20 μg/L) cadmium in blood (Cd-B $\geq$ 10 μg/L or 0.09 μmol/L) or urine (Cd-U $\geq$ 5 μg/24 h), or arsenic in urine (As-U $\geq$ 50 μg/L).[6,10–12] In 1987, the lead reporting levels were lowered to 25 μg/dL blood and 50 μg/L urine.

Cases are interviewed when the first report is received, either by telephone or in person. The interview includes questions on name and address of employer, work history (job titles and workplace exposures), industrial hygiene and exposure monitoring programs, hobbies, medical history, and demographic information. The Standard Industrial Classification (SIC) system of categorizing the employer is used.[13]

Occupational exposure is defined from the interview as a positive response to the questions of either working with or being exposed to heavy metals; in addition, if other workers from the same company have been reported to the Registry, the employer is interviewed to determine if the case could be exposed. Nonoccupational exposures include pica in the institutionalized, poisonings, avocational use of firing ranges, and home manufacture of lead shot and sinkers.

Clinical laboratories which analyze lead in state residents must be certified through proficiency testing, whether or not the laboratory is located in the state. Clinical laboratories in addition to those listed on the reporting forms have been found through a state survey of companies which store, use, or discharge toxic chemicals.[14] Employers known to be using heavy metals are interviewed to determine the number of workers using heavy metals, whether workplace air or biological samples are monitored, and which clinical laboratory is used.

Employers of reported cases are offered industrial hygiene consultation. The private physician is contacted when high levels are reported to ensure that information concerning the employer has been received and, if necessary, to assist in contacts with the employer. Companies whose employees have especially high or persistently elevated levels may be investigated by the health department or reported to OSHA.

Results

From 1982–86, 2,862 of the 3,309 cases (86.5 percent) reported to the Registry were interviewed; of these, 2,798 (97.8 percent) provided employer information. The employer could be determined from other sources such as the labora-

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TABLE 1—Number of Cases and Companies by Industrial Sector and Metal, New York State Heavy Metals Registry, 1982–86 (occupational exposures only)

	Number of Cases				Number of Companies			
Industrial Sector	Pb	Hg	As	Cd	Pb	Hg	As	Cd
Manufacturing	1241	873	52	22	98	21	3	5
Non-manufacturing	368	127	34	4	149	28	15	4
Total	1609	1000	86	26	247	47	18	9

Pb = lead; Hg = mercury; As = arsenic; Cd = cadmium. Manufacturing sector = SIC major group 20-39; non-manufacturing sector = all other SIC codes (Standard Industrial Classification).[13]

tory report in 375 of the 448 cases (83.7 percent) who were not interviewed; thus, the current employer was known for 3,173 cases (95.9 percent of total). There were no differences in metals levels between cases whose employer could and could not be determined.

A total of 2,739 occupational cases were reported from 328 companies. The cases were predominantly male (92.0 percent) and less than 40 years of age (57.1 percent); all cases were employed at their current company for at least one month. Clinical laboratories (N = 72) reported 88.9 percent of cases. Thirty-three of the clinical laboratories, including the largest reporter, were located outside the state; out-of-state sources reported 53.3 percent of cases.

Table 1 presents the distributions of reports by industrial sector. The predominant reporting was for lead (58.7 percent of cases, 75.3 percent of companies) and mercury (36.5 percent of cases, 14.3 percent of companies). The manufacturing sector (SIC major group 20–39) accounted for 80.4 percent of cases and 39.9 percent of companies. The largest numbers of non-manufacturing sector companies were reported in the service sector (SIC major group 70–89; 58 companies, 23.5 percent of total) and in the construction sector (SIC major group 15–17; 42 companies, 17.0 percent of total).

The four-digit SIC codes of industries with 20 or more cases are presented in Table 2. This criterion is arbitrary and does not imply that significant hazards do not exist in industries with fewer cases. The largest fraction of lead cases was employed in secondary smelting and refining of nonferrous metals; of mercury cases, in the manufacture of alkalies and chlorine. Industrial instrument manufacture was notable for a large number of cases and a high median urinary mercury.

Companies reported to the Registry for lead were com-

TABLE 2—Median and Maximum Heavy Metals Levels in Industries with 20 or More Cases, by Metal and Four-Digit SIC Code, New York State Heavy Metals Registry, 1982–86 (occupational exposures only)

Four-digit SIC Code and Metal	Industry Description	Reported No. of Cases	Metals levels	
			Median	Maximum
Blood Lead			(μg/dL)	(μg/dL)
1622	Bridge, tunnel, and elevated highway construction	50	56	178
1721	Painting, paper hanging, and decorating	22	50	145
2816	Inorganic pigments	37	60	87
2819	Industrial inorganic chemicals, n.e.c.	21	45	70
2899	Chemicals and chemical preparations, n.e.c.	54	47	78
3231	Glass products, made of purchased glass	110	48	340
3341	Secondary smelting and refining of nonferrous metals	343	50	88
3362	Brass, bronze, copper, copper base alloy foundries (castings)	34	53	78
3432	Plumbing fixture fittings and trim (brass goods)	58	52	89
3494	Valves and pipe fittings, except plumbers' brass goods	101	49	83
3691	Storage batteries	146	45	65
3692	Primary batteries, dry and wet	43	51	74
3714	Motor vehicle parts and accessories	74	48	78
5084	Industrial machinery and equipment	37	52	85
5093	Scrap and waste metals	26	51	94
7539	Automotive repair shops, n.e.c.	50	47	79
Urinary Mercury			(μg/L)	(μg/L)
2812	Alkalies and chlorine	494	76	916
3497	Metal foil and leaf	39	34	363
3823	Industrial instruments for measurement, display, and control of process variables	213	145	889
3829	Measuring and controlling devices, n.e.c.	33	68	309
Urinary Arsenic			(μg/L)	(μg/L)
3861	Photographic equipment and supplies	49	80	524

n.e.c. = not elsewhere classified. SIC = Standard Industrial Classification.[13] Reporting criteria: Blood lead ≥40 μg/dL or 1.93 μmol/L, urinary mercury ≥20 μg/L, urinary arsenic ≥50 μg/L.

pared with those inspected by OSHA from 1980–85 for violations of the federal lead standard. OSHA inspections were focused on the manufacturing sector to a much greater degree than Registry reports: 130 of 157 companies (82.8 percent) inspected by OSHA were in the manufacturing sector, as compared to 98 of 247 companies (39.7 percent) reported to the Registry. OSHA inspected 18 of 98 (18.4 percent) of companies reported to the Registry from the manufacturing sector, but only one of 149 (0.6 percent) companies outside the manufacturing sector.

Discussion

The Registry has several advantages compared to other surveillance methods. It is inexpensive to operate because existing data sources are used. Apart from error in the analysis of samples, there is little uncertainty regarding exposure because cases are defined by increased biological absorption. Clinical laboratory-based reporting eliminates the need to enlist the cooperation and time of individual physicians and health facilities, whose reporting is redundant to the extent that samples are sent to clinical laboratories for analysis.

Perhaps most importantly, surveillance can be carried out in industrial sectors which have a low priority for regulatory agencies and for metals for which no governmental standards exist. OSHA inspections target private companies in the manufacturing sector, larger companies, and industries with previously reported exposures. Of the four heavy metals on which Registry data are collected, a federal standard has been promulgated for lead only.

For most companies, the source of exposure is known. Lead exposure in secondary smelting and refining of nonferrous metals (SIC 3341) and storage battery manufacture (SIC 3691) results from melting of lead alloys and dust exposure during foundry or fabrication work. Alkali and chlorine manufacture workers (SIC 2812) are employed in a chloralkali plant where mercury is used as a cathode in the electrolysis of salt to produce sodium hydroxide and chlorine gas; exposure occurs when the mercury cells are opened for maintenance. Mercury exposure in industrial instrument manufacture (SIC 3823) occurs during manual assembly and fabrication of thermometers. In a small fraction of industries, heavy metals exposures have not been previously documented. The source of exposure is apparent in some, but in others more detailed investigations are necessary.

In summary, the Registry is effective in detecting companies with heavy metals exposures, particularly those outside the manufacturing sector which are not detected by OSHA inspections. Since 1982, heavy metals registries have been established in New Jersey, California, Texas, and Maryland. We recommend adoption of this surveillance method by other state health departments, and evaluation of the feasibility of a national heavy metals absorption registry.

ACKNOWLEDGMENTS

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Call for Papers: Nursing Research

The West Virginia Nurses' Association is planning a national research symposium—"Vision of Excellence: The Decade of the Nineties"—to be held at Greenbrier Hotel, White Sulphur Springs, WV, November 9–11, 1990. This two-day symposium will cover a wide range of health-related topics. The deadline for abstracts is April 15, 1990. Selected papers will be published in the conference proceedings. For further information, contact Dr. Janet F. Wang, School of Nursing, West Virginia University, Morgantown, WV 26506; telephone: (304) 293-4297.

Lead-Induced Anemia: Dose-Response Relationships and Evidence for a Threshold

Joel Schwartz, PhD, Philip J. Landrigan MD, MSc, Edward L. Baker, Jr., MD, MPH,
Walter A. Orenstein, MD, and Ian H. von Lindern, PE, PhD

Abstract: We conducted a cross-sectional epidemiologic study to assess the association between blood lead level and hematocrit in 579 one to five year-old children living near a primary lead smelter in 1974. Blood lead levels ranged from 0.53 to 7.91 μmol/L (11 to 164 μg/dl). To predict hematocrit as a function of blood lead level and age, we derived non-linear regression models and fit percentile curves. We used logistic regression to predict the probability of hematocrit values less than 35 per cent. We found a strong non-linear, dose-response relationship between blood lead level and hematocrit. This relationship was influenced by age, but (in this age group) not by sex; the effect was strongest in youngest children. In one year-olds, the age group most severely affected, the risk of an hematocrit value below 35 percent was 2 percent above background at blood lead levels between 0.97 and 1.88 μmol/L (20 and 39 μg/dl), 18 percent above background at lead levels of 1.93 to 2.85 μmol/L (40 to 59 μg/dl), and 40 percent above background at lead levels of 2.9 μmol/L (60 μg/dl) and greater; background was defined as a blood lead level below 1.88 μmol/L (20 μg/dl). This effect appeared independent of iron deficiency. These findings suggest that blood lead levels close to the currently recommended limit value of 1.21 μmol/L (25 μg/dl) are associated with dose-related depression of hematocrit in young children. (*Am J Public Health* 1990; 80:165–168.)

Introduction

Lead-induced anemia is produced principally by two mechanisms: impairment in heme biosynthesis, and increased rate of red blood cell destruction.[1]

Lead-induced anemia has been reported to occur in children at a blood lead level of 1.93 μmol/L (40 μg/dl).[2]* However, no clear dose-response relationships between blood lead level and hematocrit have been established. Further, almost no data are available to assess whether this relationship extends downward to blood lead levels below 1.93 μmol/L (40 μg/dl), to ascertain the fraction of children with lead-induced anemia at various blood lead levels, or to determine whether there exists a threshold, below which the association is no longer evident.

To assess dose-response relationships between blood lead level and hematocrit, we examined data obtained in 1974 during an epidemiologic study of lead exposure in children living near a lead smelter.[3]

Methods

A cross-sectional prevalence study of lead exposure in children living near a lead ore smelter in Kellogg, Idaho, conducted in August 1974, found blood lead levels of 1.93 μmol/L (40 μg/dl) and above in 285 (41.9 percent) of 919 children. Within 1.6 km of the smelter, 170 (98.8 percent) of 172 children had blood lead levels above 1.93 μmol/L (40 μg/dl), the highest being 7.91 μmol/L (164 μg/dl). Blood lead levels were positively associated with exposures to airborne lead (correlation coefficient, $r = 0.72$; $p < 0.0001$) as well as with exposures to lead in soil ($r = 0.5$; $p < 0.0001$).[4]

Blood samples (drawn in lead-free vacuum tubes) were analyzed for hematocrit by centrifugation and for erythrocyte protoporphyrin (EP) by the method of Granick, *et al.*[5] Samples were analyzed for lead by modified Delves' cup atomic absorption spectrophotometry.[6] These methods have previously been described in detail.[3]

A total of 1,058 children, ages 1 to 9 years, were identified for whom data on both blood lead and hematocrit were available. In the present analysis, we focus on the 579 children in the age group 1 to 5 years, since children in this age range appear most sensitive to the hematotoxic effects of lead.

Anemia has been defined variously, with the lower bound of the normal range of hematocrit values in young children varying from 33 to 35 per cent.[7–10] We defined anemia as an hematocrit value less than 35 percent. To assess the relation between blood lead level and hematocrit in children using this definition, we fit logistic regressions, which predict the probability of anemia at different age, sex and blood lead levels.[11]

Initial examination of a raw scatter plot of hematocrit versus blood lead level showed a non-linear dose-response relationship. There was little variation in hematocrit at low blood lead levels and a negative relation with lead at higher levels.

Rather than fitting this relationship directly in a non-linear parametric regression, which imposes an explicit functional form on the data, we employed non-linear smoothing techniques developed for robust graphical analysis.[12,13]

Results

The smoothed plot of the data follows a logistic-shaped curve (Figure 1), which is relatively flat at blood lead levels below 0.97 μmol/L, becomes increasingly steep as blood lead levels rise, and, at very high lead levels, flattens again. This curve can be represented by the logistic function:

$$\text{Hct} = A/(1 + \text{EXP}(B_o + B_1{*}\text{lead} + B_2{*}\text{age})) + C \qquad (1)$$

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*1 μmol/L = 20.72 mg/dl blood lead.

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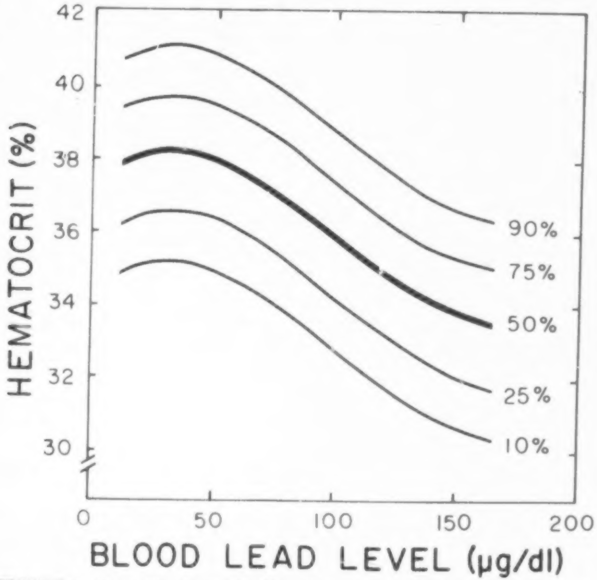

FIGURE 1—Percentile Curves for Hematocrit versus Blood Lead Level in 1–5 Year Old Children, Northern Idaho, 1974

The logistic regression analysis showed a strong association between elevation of blood lead level and the probability of anemia (regression coefficient = 0.3083; (SE = 0.0061). Year of age was also a significant predictor of low hematocrit (coefficient = −0.3831; SE 0.1134). In this age group, sex was not a predictor of anemia. The predicted probability of anemia in relation to blood lead level is plotted graphically for each year of age in Figure 2.

To illustrate the fit of the predicted curve to the actual data, Figure 3 shows both the predicted curve and actual points for the mean proportion of children with anemia for each interval of 0.48 μmol/L (10 μg/dl) in blood lead level after controlling for age. The highest point represents the mean for the five children with blood lead levels greater than 5.79 μmol/L (120 μg/dl).

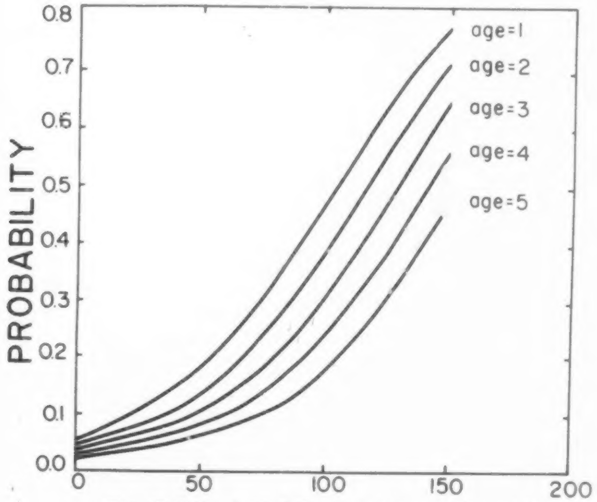

FIGURE 2—Predicted Probability of Hematocrit <35% by Age and Blood Lead Level in 1–5 Year Old Children, Northern Idaho, 1974

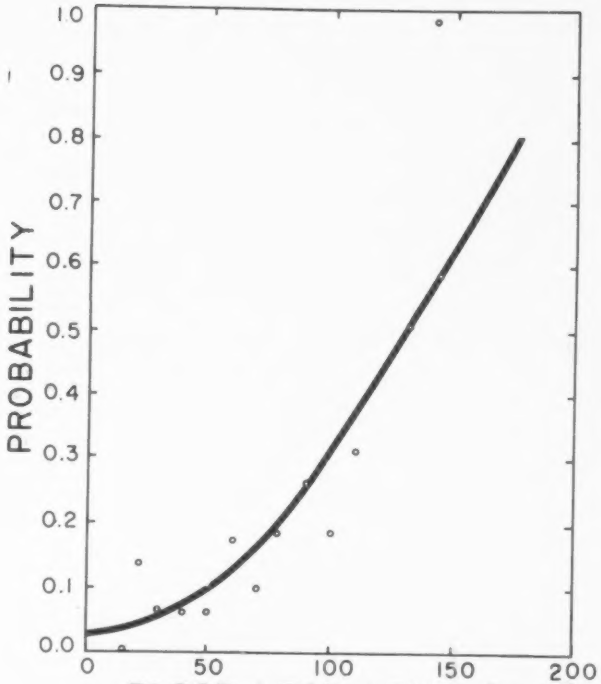

FIGURE 3—Mean Probability of Hematocrit <35% by Blood Lead Level, Northern Idaho, 1974 (This line is the fitted logistic curve for the probability of hematocrit <35% after adjusting, by regression, to the mean age)

Since Figure 2 suggests that the increased risk of depressed hematocrit begins at lower blood lead levels in one year-olds than in five year-olds, we calculated the mean proportion of children in each year of age with an hematocrit below 35 percent at various blood lead levels (Table 1). To compensate for the reduced sample size, we used increments of 0.97 μmol/L (20 μg/dl) in blood lead levels. While the sample size was too small to draw strong conclusions, the results were concordant with the regression analysis (Table 1). In one year-old children the probability of anemia was 0 percent at blood lead levels below 0.97 μmol/L (20 μg/dl), 2 percent at levels between 0.97 and 1.93 μmol/L (20 and 39 μg/dl), 18 percent at levels between 1.93 and 2.90 μmol/L (40 and 60 μg/dl), and 40 percent at blood lead levels above 2.90 μmol/L (60 μg/dl).

We also examined actual hematocrit values rather than proportion anemic. In Figure 1, the 20th, 25th, 75th, and 90th percentile curves reflect the broad variation in hematocrit levels due to age, iron deficiency and other factors. Percentile curves were estimated by confirming the homoscedasticity of the residuals from the smooth curve and then estimating their standard deviation from their interquartile range; percentiles were then calculated from the standard deviation. Table 2 presents the results of nonlinear regression of hematocrit on age and on blood lead level.

Figure 4 plots predicted hematocrit levels by age and blood lead level. Again, it shows the strong effect of age on hematocrit in young children, and demonstrates the interrelationship between age and blood lead level.

TABLE 1—Distribution of Population of Children with Anemia by Year of Age and Blood Lead Level, Northern Idaho—1974

	Blood Lead Level μmol/L*									
Age (Yrs)	0.–0.92		0.93–1.88		1.89–2.85		2.86 +		Total	
	# Anemic	Total	# Anemic	Total	# Anemic	Total	# Anemic	Total	# Anemic	Total
1	0	6	1	47	6	33	6	15	13	101
2	1	3	5	49	6	35	6	18	18	105
3	2	7	3	56	2	33	3	29	10	125
4	1	1	4	60	1	23	4	24	10	108
5	0	3	2	78	1	37	1	22	4	140
Total	4	20	15	290	16	161	20	108	55	579

*1 μmol/L = 20.72 μg dl whole blood lead level.

TABLE 2—Non-Linear Regression Model for Hematocrit versus Blood Lead Level, 1–5 Year-Old Children, Northern Idaho, 1974*

Effect	Parameter	Estimate	Standard Error	p-Value
Constant	A	39.54	0.79	0.0001
Constant	B_o	−3.112	0.446	0.0001
Blood Lead Level	B_1	0.0133	0.0041	0.0005
Age	B_2	−0.2016	0.0905	0.0129

*The parameters A, C, and B_o all relate to the zero lead intercept and were not simultaneously significant. C was therefore deleted (reducing the error of sum squares) to give the model shown in this table.

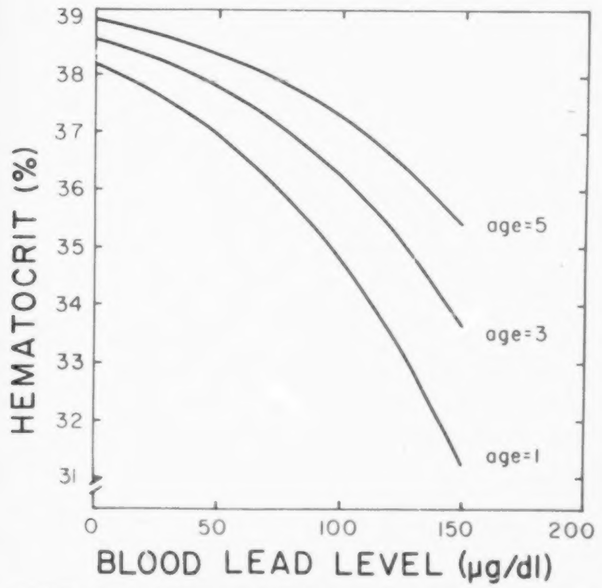

FIGURE 4—Predicted Hematocrit Value versus Blood Lead Level, by Age, in 1-5 Year Old Children, Northern Idaho, 1974

Discussion

The results of this study show a strongly negative, non-linear dose-response relationship between blood lead level and hematocrit in children, strongest in the youngest children. Although there is variability among children in the effect of lead on hematocrit, the data from this study suggest that lead-induced anemia is a biologically and clinically important consequence of lead absorption, even at low levels of exposure.

Iron deficiency may have accounted for some of the observed effect of lead on hematocrit since the effects of lead and iron deficiency on hematocrit have been shown to be strongly interactive.[14] Although data on the iron status of the children in our study were not obtained, the strong correlation between low socioeconomic status, elevated lead levels, and iron deficiency typically seen in urban populations did not exist in our study population. In our population, blood lead levels were determined principally by distance from the smelter and were not correlated with socioeconomic status,[3,4] hence lead exposure and iron deficiency were probably not correlated in these data.

The smoothed dose-response curve of hematocrit versus blood lead level (Figure 1) shows a slight, paradoxical increase in hematocrit at low blood lead levels, followed by a decrease, whose slope steepens until blood lead levels reach 4.83 μmol/L (100 μg/dl) and then subsequently flattens out again. The slight increase in hematocrit observed at the lowest blood lead levels may be artifactual. Since children with lowest blood lead levels came from a rural valley distant from the smelter town, with low family income and socioeconomic status, nutritional deficiency may have been more common there and account for the observed results. Alternatively, the effect may be due to differential partitioning of lead among the soft tissues of the body. At equilibrium, the fraction of lead stored in any body compartment will be proportional to the size of the compartment. At body burdens too low for lead to affect hemoglobin levels, higher hemoglobin levels could increase the fraction of lead stored in red blood cells thus causing the positive correlation seen here.

These data imply that there is no margin of safety between the threshold blood lead level at which lead begins to cause depression of hematocrit in young children and the blood lead level currently considered by the Centers for Disease Control to represent the upper limit of acceptability—1.21 μmol/L (25 μg/dl)[15] These findings complement other data showing deleterious effects of lead on heme synthesis of blood lead levels of 0.48 to 0.97 μmol/L (10–20 μg/dl), on vitamin D metabolism at blood lead levels below 1.21 μmol/L (25 μg/dl), on peripheral nerve conduction velocity at 0.97 to 1.45 μmol/L (20 to 30 μg/dl) and on cognitive function at 0.72 to 0.97 μmol/L (15 to 20 μg/dl).[16-22] Taken together, these findings suggest that the current recommendation on the acceptable upper limit for blood lead level in children requires reexamination.

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Call for Applications: Minority Medical Faculty Development Program 1990

The Robert Wood Johnson Foundation's Minority Medical Faculty Development Program offers four-year, postdoctoral research fellowships to minority physicians who have demonstrated superior academic and clinical skills and who are committed to careers in academic medicine and biomedical research. Each of the Fellows selected (up to eight) will receive an annual stipend of up to $50,000, complemented by a $25,000 annual grant toward support of research activities. Each Fellow will study and conduct research under the supervision of a senior faculty member located at an academic center of biomedical research noted for the training of young faculty and pursuing lines of investigation that are of interest to the Fellow. Priority will be given to basic research. It is expected that Fellows will be spending no less than 70 percent of their time in pursuit of research activities.

The Minority Medical Faculty Development Program seeks to increase the number of minority faculty who can encourage and foster the development of succeeding classes of minority physicians. At present, only 1,281 (3.2 percent) of the nation's 40,136 full-time medical faculty members are minorities.

Foundation support for the four-year fellowships will be provided initially for *two years*, with funding for the third and fourth years dependent upon review and recommendation by the National Advisory Committee. The fourth year of support may be used to assist the Fellow in his or her first year as a full-time medical faculty member.

To be eligible, applicants must be minority physicians who are US citizens, have excelled in their education, are now completing or will have completed formal clinical training (internship, residency, or clinical fellowship), and are committed to academic careers.

April 2, 1990—Deadline for request of application materials.

By April 30, 1990—Receipt of completed applications.

July 1990—Interviews with National Advisory Committee and consultants in biomedical research.

December 1990—Official notification of Fellows.

Requests for applications and inquiries should be addressed to: Harold Amos, PhD, Program Director, Minority Medical Faculty Development Program, P.O. Box 446, Brookline, MA 02147. Tel: 617/732-1947.

An Analysis of Occupational Risks for Brain Cancer

Ross C. Brownson, PhD, John S. Reif, DVM, MSc, Jian C. Chang, PhD, and James R. Davis, MS

Abstract: We evaluated the risks of brain cancer in relation to employment history in a case-control study of 312 cases and 1,248 cancer controls. Subjects were identified through the Missouri Cancer Registry for the period 1984 through 1988. Job classification was based on data routinely abstracted from hospital records. Elevated risks were identified for certain white collar occupations: for men employed in engineering, the odds ratio (OR) = 2.1; 95% confidence interval (CI) = 0.4, 10.3; for social science professionals, the OR = 6.1; 95% CI = 1.5, 26.1. Among occupations with potential exposure to occupational carcinogens, increased risks were observed for men employed in agricultural crop production (OR = 1.5; 95% CI = 1.0, 2.4), printing and publishing (OR = 2.8; 95% CI = 1.0, 8.3), and brickmasons and tilesetters (OR = 2.5; 95% CI = 0.5, 11.5). Most of elevated brain cancer risks were due to astrocytic cancers, but the excess among agricultural workers occurred in other cell types. No increase in risk was noted for current cigarette smokers (OR = 0.9; 95% CI = 0.7, 1.5) or ex-smokers (OR = 1.0; 95% CI = 0.7, 1.5). This exploratory study indicates a need for further studies of occupational risks of brain cancer. (*Am J Public Health* 1990; 80:169–172.)

Introduction

Cancer of the brain has been the focus of relatively few epidemiologic studies. Increased risks have been noted for several white collar occupations including accountants,[1] sales managers,[2] engineers,[2,3] teachers,[1] lawyers and judges,[2] and postal clerks.[3] Among occupations with potential carcinogenic exposures on the job, excess brain cancer risk has been suggested for workers employed in the rubber industry,[4,5] polyvinyl chloride production,[6-9] farming,[10-13] electrical industries,[14-16] pharmaceutical production,[17] and production of petroleum and petrochemicals.[18,19] Cigarette smoking has also been linked to brain cancer risk.[20]

Most previous studies of occupation and brain cancer have relied on mortality data, with concomitant potential for misclassification of brain cancer on death certificates.[21] Few previous studies have evaluated differences in occupational risks by histologic type of brain cancer. The current study was based on data from a statewide cancer registry.

Methods

Cases and controls were identified through the Missouri Cancer Registry, maintained by the Missouri Department of Health. The Registry has been collecting data on cancer cases from public and private hospitals since 1972, with reporting mandated by law since 1984.

The case group consisted of White males diagnosed with histologically confirmed brain and other central nervous system cancer (*International Classification of Diseases for Oncology* [ICD-O] codes 191 and 192)[22] between January 1984 and December 1988. Selection was limited to White males due to the small number of non-Whites and the lack of reported occupational diversity among females. The mean age among cases was 54.6 years.

A frequency-matched sample of controls was chosen from all other white male patients diagnosed with cancers in the same time period, excluding cancers of ill-defined or unknown primary sites (ICD-O site codes 195 and 199). For each case, four controls were randomly selected within each of six age strata. The mean age of controls was 55.6 years. The analyses included a total of 312 cases and 1,248 controls. The distribution of cancer sites among controls is shown in Table 1.

Data on occupation and tobacco use are routinely abstracted from the hospital medical record at the time of diagnosis and submitted to the Missouri Cancer Registry as part of cancer incidence reporting. Hospital registrars record information for cancer patients using a standardized protocol and the occupational data are subsequently coded by trained cancer registry specialists using the three-digit 1980 US Census codes.[23] The occupational history reported to the Registry is intended to represent the usual or longest held job of the cancer patient. Smoking history for each patient is characterized according to status (i.e., never, former, or current use) and, among current smokers, according to the number of packs smoked per day.

The present analysis was based on subjects who had codable occupational information (i.e., either occupation or industry or both variables). The remaining subjects had no occupational information or noninformative job titles such as "retired." Duration of employment was only occasionally recorded and not used in these analyses. Of the initially eligible subjects, 34 percent of potential cases and 38 percent of potential controls were excluded due to missing occupational information.

The odds ratio (OR) was the measure of association between occupation and brain cancer. Gart's method[24] was used to calculate maximum likelihood estimates of the OR

TABLE 1—Distribution of Cancer Sites in Control Group, Missouri, 1984–88

Site of Cancer	Number	Percent
Lip, oral cavity, and pharynx	57	5
Digestive organs and peritoneum	276	22
Respiratory system and intrathoracic organs	356	29
Skin	21	2
Bones and connective tissue	24	2
Genitourinary organs	305	24
Leukemia	35	3
Lymphoma and multiple myeloma	63	5
Other sites	111	9
Total	1,248	100

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and corresponding 95 percent confidence intervals (CI). For each occupation and industry with two or more cases, age- and smoking-adjusted ORs were computed. The reference category consisted of all persons not in the occupation of interest.

Separate analyses were conducted for the astrocytic series (ICD-O morphology codes 93803, 93823, 94003-94413) and for "other" cell types of brain cancer. The astrocytic series was comprised primarily of unspecified astrocytomas (31 percent), unspecified glioblastomas (51 percent), and anaplastic astrocytomas (7 percent). "Other" cell types included oligodendrogliomas (31 percent), unspecified cell types (20 percent), and unspecified ependymomas (14 percent). All controls were used in each cell type analysis.

For analyses by smoking history, we excluded controls with cancers of smoking-related sites (lip, oral cavity, esophagus, larynx, lung, and bladder—ICD-O site codes 140-141, 143-145, 150, 161, 162, and 188).

Results

Risk of brain cancer according to industry of employment is shown in Table 2. Men employed in agricultural crop production had a moderately elevated risk (OR = 1.5; 95% CI = 1.0, 2.4). An OR of 2.8 (95% CI = 1.0, 8.3) was observed for workers in printing and publishing. Several other industries were associated with risks of 2.0 or greater, although estimates were based on small numbers. These included men employed in miscellaneous repair services (OR = 2.6; 95% CI = 0.5, 13.0), engineering services (OR = 2.1; 95% CI = 0.4, 10.3), and justice, public order, and safety (OR = 2.1; 95% CI = 0.9, 4.8).

As shown in Table 3, excess risks were found for social science professionals (OR = 6.1; 95% CI = 1.5, 26.1) and for police and fire protection services (OR = 2.2; 95% CI = 1.0, 4.7). The excess among police and fire protection services included an OR of 2.9 (95% CI = 0.8, 11.0) for police and detectives and an OR of 2.0 (95% CI = 0.4, 9.6) for firefighters. An OR of 2.5 (95% CI = 0.5, 11.5) was observed for brickmasons and tilesetters, although there were only three cases in that job category.

TABLE 2—Age- and Smoking-Adjusted Odds Ratios* (OR) and 95% Confidence Intervals (CI) for Brain Cancer by Industry, Missouri, 1984–87

US Census Codes	Industry	Number		OR (95% CI)
		Cases	Controls	
010	Agriculture—crop production	22	62	1.5 (1.0,2.4)
060	Construction	26	135	0.8 (0.5,1.3)
100–122	Food manufacturing	6	36	0.7 (0.3,1.7)
171–177	Printing and publishing	7	10	2.8 (1.0,8.3)
180–192	Chemical manufacturing	4	14	1.1 (0.3,3.6)
270–301	Metal manufacturing	7	22	1.3 (0.5,3.2)
351–370	Transportation equipment manufacturing	20	83	0.9 (0.6,1.6)
391–392	Miscellaneous and unspecified manufacturing	9	46	0.8 (0.4,1.7)
400–432	Transportation	31	121	1.1 (0.7,1.7)
440–442	Communications	5	15	1.4 (0.5,4.1)
460–472	Utilities and sanitary services	3	26	0.5 (0.1,1.7)
591	Department stores	4	12	1.3 (0.4,4.5)
601	Grocery stores	5	12	1.6 (0.5,5.1)
641	Eating and drinking places	2	6	1.6 (0.2,8.9)
700–712	Finance, insurance, real estate	10	38	1.0 (0.5,2.2)
751	Automotive repair shops	11	28	1.7 (0.8,3.7)
760	Miscellaneous repair services	3	5	2.6 (0.5,13.0)
842, 850	Schools	13	37	1.3 (0.6,2.5)
880	Religious organizations	3	9	1.0 (0.2,4.1)
882	Engineering services	3	5	2.1 (0.4,10.3)
910	Justice, public order, safety	10	19	2.1 (0.9,4.8)
999	Industry not reported	21	93	0.9 (0.5,1.5)

*Odds ratios are reported for job categories that contained two or more cases.

TABLE 3—Age- and Smoking-Adjusted Odds Ratios* (OR) and 95% Confidence Intervals (CI) for Brain Cancer by Occupation, Missouri, 1984–87

US Census Codes	Occupation	Number		OR (95% CI)
		Cases	Controls	
003–037	Managers, administrators, executives	34	146	0.9 (0.6,1.3)
043, 063–083	Science professionals	2	11	0.6 (0.1,3.2)
044–059	Engineers	7	20	1.2 (0.5,3.2)
084–089	Physicians and dentists	3	13	0.8 (0.2,3.2)
113–159	Teachers	6	26	0.8 (0.3,2.1)
163–175, 183–184, 195–197	Social science professionals	6	4	6.1 (1.5,26.1)
176–177	Clergy	2	9	0.7 (0.1,3.3)
178–179	Lawyers, judges	3	10	1.1 (0.2,4.5)
185–194, 198–199	Artists, entertainers, athletes	3	9	1.2 (0.3,5.1)
213–235	Technicians, except health	11	27	1.6 (0.8,3.5)
243–285	Sales	25	90	1.1 (0.7,1.8)
303–389	Office supervisors and workers	16	43	1.4 (0.7,2.6)
413–427	Police, fire protection services	12	22	2.2 (1.0,4.7)
448–454	Building services	6	28	0.9 (0.3,2.3)
473–489	Farming	21	80	1.1 (0.6,1.7)
503–549, 864	Mechanics and repairers	22	68	1.4 (0.8,2.4)
553, 563–565	Brickmasons and tilesetters	3	6	2.5 (0.5,11.5)
554, 567–569	Carpenters	6	43	0.6 (0.2,1.5)
557, 585–587	Plumbers and pipefitters	3	12	1.2 (0.3,4.5)
558, 566, 588–589, 599, 863, 865, 866, 869	Construction, not elsewhere classified	11	36	1.3 (0.6,2.7)
594, 843–859	Pavers, surfacers, material movers	4	14	1.2 (0.3,3.9)
633–639, 644–655, 689–693, 796–799, 873	Precision production/tool and die makers, machinists, sheet metal workers	12	43	1.2 (0.6,2.4)
686–688	Food production	3	8	1.6 (0.3,6.8)
753–779	Machine operators, miscellaneous	12	72	0.7 (0.4,1.4)
785	Assemblers	5	15	1.4 (0.4,4.2)
803–814	Motor vehicle operators	22	83	1.2 (0.7,2.0)
889	Laborers, except construction	7	46	0.6 (0.2,1.4)
905–999	Unemployed/last job armed forces/unknown	24	119	0.8 (0.5,1.4)

*Odds ratios are reported for job categories that contained two or more cases.

Evaluation of risks according to histologic type of brain cancer showed that the elevated risk in agricultural workers was confined to "other" cell types (OR = 5.3; 95% CI = 1.9, 14.3). The elevated overall risk for workers in printing and publishing was equally distributed between astrocytic and "other" cell types. Elevated risks of astrocytomas were observed for automobile repair work (OR = 1.9; 95% CI = 1.0, 4.1) and justice, public order, and safety (OR = 2.4; 95% CI = 1.0, 5.6).

Social science professionals showed elevated risks for both astrocytomas (OR = 4.7; 95% CI = 1.0, 22.9) and "other" cell types (OR = 19.0; 95% CI = 2.1, 145.6). The excess risk among police and fire protection workers was confined to the astrocytic cell series (OR = 2.3; 95% CI = 1.0, 5.1). Farmers had an increased risk for "other" cell types (OR = 3.7; 95% CI = 1.4, 9.8). Machinists had an OR of 1.8 (95% CI = 0.6, 5.2) for astrocytomas.

There were 33.7 percent current smokers and 19.9 percent ex-smokers among cases, compared with 34.4 percent current smokers and 18.5 percent ex-smokers among controls. Current smokers had an OR of 0.9 (95% CI = 0.7, 1.3) and ex-smokers had an OR of 1.0 (95% CI = 0.7, 1.5).

Discussion

The present study confirmed previously reported occupational associations[2,3,25] between brain cancer risk and employment in certain white collar occupations. These increased risks may be related in part to better access and utilization of medical services and, hence, a greater likelihood of brain cancer diagnosis.[26]

An increased risk of brain cancer among agricultural workers has been reported in a number of recent studies,[10–13,27] although several others have failed to identify such a relationship.[21,28,29] In the current study, the elevated risk among farmers was confined to "other" cell types. Mussico, *et al*,[30] recently found an increased risk of brain gliomas in relation to use of insecticides or fungicides. Animal studies[31] and elevated levels of organochlorine compounds in adipose tissue of glioblastoma patients[32] adds biological plausibility to the hypothesis that agricultural chemical exposure may be partially responsible for the excess risk of brain cancer in farmers.

The increased risk seen in this study among workers in the printing and publishing industry was unexpected, although one previous study[33] suggested an elevated brain cancer risk in commercial pressmen. Since printers are exposed to a variety of chemicals, this relationship deserves further scrutiny.

Several *a priori* suspected high-risk occupations showed indications of elevated risks in our study; however, some of these comparisons were limited by small numbers. For example, we found slight elevations in risk for chemical manufacturers, mechanics and repairers, and precision metal workers such as machinists, all of which have been reportedly previously.[1,25,29,34] We identified an elevated risk of astrocytomas for automobile mechanics, who are potentially exposed to solvents, lubricating oils, and hydrocarbons.[25] Rushton, *et al*,[35] found elevated brain cancer mortality among bus mechanics in the United Kingdom. Our data suggested an increase in risk of astrocytomas for brickmasons and tilesetters. This association was also reported in a death certificate study from Massachusetts.[34]

We found no indication of an increased risk among smokers. This contrasts with a recent study[21] but is consistent with two earlier studies.[36,37]

The current study has several limitations. We relied on occupational information collected in conjunction with cancer incidence reporting. Of the initially eligible study subjects, over one-third had missing occupational information. The Registry obtains job information based on the hospital medical record at the time of diagnosis and some degree of misclassification is likely. A recent study that compared risk factor information contained in the Missouri Cancer Registry record with information obtained by interview found an agreement of 70 percent between Registry-reported occupation and usual occupation reported by interview.[38] In addition, occupational misclassification was apparently random and would therefore bias relative risk estimates toward unity.[39]

We selected controls from other cancer patients recorded on the Registry, with the attendant advantages and disadvantages.[40,41] We compensated for the relation between smoking and other cancers in sub-selection of controls and adjustment in the analyses.

The strengths of our study included greater accuracy in hospital-reported tumor diagnosis, compared with death certificates,[21] and the ability to analyze by histologic type.

This exploratory study suggests leads for future investigations with more information on occupational exposures and other possible confounders.

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The Robert Wood Johnson Clinical Scholars Program

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Work-Related Psychosocial Stress and Risk of Preterm, Low Birthweight Delivery

CHARLES J. HOMER, MD, MPH, SHERMAN A. JAMES, PhD, AND EARL SIEGEL, MD, MPH

Abstract: We investigated whether work-related psychologic stress—defined as work characterized by both high psychologic demands and limited control over the response to these demands—increases a woman's risk of delivering a preterm, low birthweight infant.

We studied 786 employed pregnant women included in the National Longitudinal Survey of Labor Market Experience, Youth Cohort (NLSY), a nationally representative sample of 12,686 young adults. Data concerning work status, job title, and other factors affecting pregnancy outcome were obtained from the NLSY. Assessment of job experience was based on job title, using an established catalogue of occupation characteristics.

After accounting for the physical exertion entailed in a job, occupational psychologic stress as measured by job title was not associated with preterm, low birthweight delivery for the sample as a whole (Relative risk = 1.16, 95% confidence interval .45, 2.95). For those women who did not want to remain in the work force, work-related stress increased their risk of experiencing this outcome (RR = 8.1, 95% CI 1.5, 50.2). Personal motivation toward work, as well as the physical effort of work, should be considered in evaluating the impact of a job's psychologic characteristics on pregnancy outcome. (*Am J Public Health* 1990; 80:173–177.)

Introduction

A greater proportion of pregnant American women now work outside the home for pay than ever before.[1] Whether certain types of work adversely influences the outcome of pregnancy therefore has substantial import for maternal and child health.

Although a number of studies have examined the broader question of the influence of life stress on pregnancy outcome,[2] relatively little research exists in the specific area of work-related stress.[3] Fox, *et al,* compared low birthweight and perinatal mortality rates among active duty military women with a control group matched for race and parity. He found both outcomes significantly more frequent among the active duty group.[4] Mamelle, *et al,* also investigated the combined effects of psychological and physical stressors. She found that women with high levels of occupational fatigue—defined as work with a combination of prolonged standing, excessive routine, exposure to vibrating machinery, exposure to cold, heat or chemicals, and lifting—experienced higher risk of preterm delivery.[5] These studies both exhibit major limitations, including inadequate control for confounding factors and, in Mamelle's study, the possibility of recall bias.[6]

We evaluate here the impact of job related psychologic stress on pregnancy outcome for a national sample of *young* mothers.

Methods

Study Design, Population

We included in this study those women in the ongoing National Longitudinal Survey of Labor Market Experience, Youth Cohort (NLSY) who met study criteria.[10] The NLSY consists of a national probability sample of 12,686 youth who ranged in age from 14–21 years when selected in 1979. The survey intentionally oversampled Blacks, Hispanics, and economically disadvantaged Whites. The use of NLSY sample weights in analyses corrects for this oversampling and allows for generalization to the entire US population for this age group. All analyses presented here incorporate sample weights.

Criteria for inclusion in the study population included: female sex, having had a singleton live birth prior to the 1983 interview, and having worked during the most recent pregnancy. Additional technical criteria necessary for including an NLSY respondent in the analyses were availability of a US Census Bureau occupational and industry code for the job held during pregnancy, of job characteristic information for that job code, and of birth outcome data—birthweight and gestation—for the particular pregnancy (see below). Of the initial 12,686 NLSY respondents, 772 women met all criteria (see Appendix).

Pregnancy Outcome and Risk Factors

The survey respondents provided detailed information concerning family background, social, and educational factors, as well as a thorough job history, at the initial interview in 1979. Since that time, the survey youth have answered an extensive in-person questionnaire annually. Women respondents first gave information concerning their reproductive experiences in 1982, retrospectively describing pregnancy experiences in all previous years. The 1983 interview updated this information and included the weight and number of weeks early or late for the most recent live birth. Only the most recent pregnancy was used for these analyses. The NLSY sample diminished less than 5 percent from 1979 through 1983, with little variation among race or sex groupings.[11]

Using the interview data available from 1979 through 1983, we ascertained information concerning many of the factors known or suspected to influence risk of preterm or low birthweight delivery.[2] Specifically, we abstracted information concerning maternal race (Black, Hispanic, other), maternal age at the time of the infant's birth, maternal federal poverty status during the year of infant birth, maximum grade completed by the mother prior to the infant's birth, grade completed by the infant's maternal grandmother as of 1979, maternal marital status at the time of the infant's birth, smoking during pregnancy, alcohol use during pregnancy,

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month began prenatal care, height, prepregnancy weight, weight gain during pregnancy, prior induced abortions, prior total pregnancy failures, parity, and prior low birthweight delivery. Fertility and obstetric data were obtained from the 1982 or 1983 interviews; sociodemographic data were obtained in most cases from the interview most closely following the infant's birth.

Job Characteristics Data

The yearly NLSY interview obtains a detailed job history, including stop and start dates for all jobs. The NLSY staff provide a US Census Bureau occupation and industry code for each job worked for 20 hours per week for longer than nine weeks. Through the use of this history, the child's birth date, and the estimated gestation, we identified those jobs held during pregnancy. Job codes were available for some women working at a lesser level of intensity during pregnancy, provided that the woman worked over 20 hours for nine weeks in the same job prior to the pregnancy. We also derived a measure of the intensity of work during pregnancy, which encompassed both the proportion of pregnancy worked and the hours worked per week.

We used the Job Characteristics Scoring System developed by Robert Karasek to determine job control, job demands, and work-related physical exertion.[12] This system is based on analyses of the pooled responses to the three Quality of Employment Surveys of 1969, 1972, 1977 by a large, representative national samples of workers.[13] Karasek and colleagues derived several scales by factor analysis of these survey responses. Scales were developed separately for men and women, and apply to most US Census Code Occupational/Industry codes.

High values for job control reflect both the skill and creativity required to perform a job (skill discretion), as well as the freedom to make decisions about the work (decision authority). The measure is both reliable and valid.[12]

The psychologic workload measure reflects a job's pace and pressure (e.g., job requires working very hard). Although somewhat less reliable, this measure maintains face validity.[12]

The physical exertion measure is based on one question asked in two of the three Quality of Employment Surveys, specifically, "Does your job require lots of physical effort?" The measure has both face and concurrent validity.[12,*]

Attitudes Toward Working

All respondents were asked in 1979 and 1982 about their aspirations at age 35 ("What would you like to be doing at age 35?": same job, different job, or having a family). If a woman responded "Having a family," she was then asked "would you like to work in addition to having a family?" For these analyses, we classified a woman as uncommitted toward paid work if she responded in the 1979 survey that she did not want to work at age 35 in addition to having a family.

Analysis

We defined each demand/control quadrant by dichotomizing demand and control at the median level of each among NLSY working pregnant women. We considered three related outcome measures:

- preterm, low birthweight, defined as gestation less than 38 weeks and birth weight 2500 grams or under (5.5 pounds);
- low birthweight itself, i.e., birthweight 2500 grams or under;
- birthweight viewed as a continuous value.

We compared the frequency of preterm, low birthweight delivery, frequency of low birthweight, and mean birthweight between women in job characterized by both high demand and low control (n = 262) with women in jobs with all other combinations of demand and control (n = 510).

We dichotomized the physical exertion measure into high and low exertion groups at the median value among NLSY working pregnant women for those analyses incorporating this measure.

We undertook weighted multiple logistic regression to determine the effect of job-related stress on risk of delivering a preterm/low birthweight or low birthweight infant.[14] In these regressions, we adjusted for the potentially confounding influence of: maternal race (Black, Hispanic, other), maternal age, maternal federal poverty status during the year of infant birth, maximum grade completed by infant's mother prior to birth, education of maternal grandmother of 1979, marital status at time of birth, smoking during pregnancy, alcohol use during pregnancy, month began prenatal care, prepregnancy weight, prior induced abortions, prior total pregnancy failures, parity, and prior low birth weight delivery, as well as work-related physical exertion and our measure of work intensity. Similar weighted multiple linear regression analyses were used to determine the effect of exertion on birthweight.

We explored whether the relationship between stressful work and preterm low birthweight varied among those more or less motivated towards work through stratification and through specification of interaction terms (stress × motivation) in the logistic and linear models.

We weighted all analyses with NLSY Sample Weights, which reflect the probability of selection relative to the US population for this age group, and also reflect non-response in any given year.[15] Design effect factors in the NLSY are generally small (less than 2), and are not included in the analyses presented here.**

With the sample size of 780, we could detect a relative risk associated with job stress on the likelihood of delivering a preterm, low birthweight infant of 1.8 with a power of 90 percent.

Results

Sample Attrition and Missing Data

Among women who worked during pregnancy, several differences existed between those women included in the final sample and those excluded due to missing information (either no Census job code or job characteristic score) (Table 1). Most of the exclusions were due to the absence of a job code, in turn due to these women having worked for limited hours. Excluded women were as a group socioeconomically disad-

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[*] Homer CJ, Beresford SAA, James SA, Siegel E, Wilcox S: Work-related physical exertion and risk of preterm, low birthweight delivery. Paediatric and Perinatal Epidemiology (in press).

[**] Unpublished technical documentation, Center for Human Resources Research, Ohio State University.

TABLE 1—Comparison of Study Women and Excluded Working Women

Measure	Worked, No Census Code	Worked, Coded, No JCSS Score	Final Sample	All NLSY Mothers
	n = 207	n = 54	n = 786	n = 2375
Maternal				
% Black	16.7	27.3*	15.7	21.9
% Hispanic	10.7	5.7	7.3	8.9
Height (in)	64.0	64.0	64.1	63.9
Weight (lbs)	128.6	122.0	126.9	127.0
Weight gain (lbs)	31.7	29.7	31.6	31.0
Education (years)	11.0*	12.0	11.9	11.2
Prior pregnancy failures (mean)	.15	.09	.15	.13
% Prior LBW	3.3	3.9	2.4	4.0
% Less than federal poverty level	36.9*	36.4	15.9	29.0
% First trimester prenatal care	74.8*	82.0	84.5	80.0
% Smoked	50.3*	32.2	40.1	43.0
Infant				
Birthweight (grams)	3245	3145	3295	3304
% LBW	7.1	5.1	6.25	8.3
% Preterm/LBW	4.3	1.7	3.4	4.7

*Excluded group differs from final sample at $p < .05$, by chi-square or T-test

vantaged relative to other working women, and were more similar to the population of non-working women.

Validation of Measures

The mean birthweight of the infants in the NLSY was 3304 grams. Preterm infants were smaller than term infants, with a mean weight of 2697 grams versus 3389 grams for term infants. Of the 2375 infants included in this study, 8.3 percent weighed less than 2500 grams. Just under 5 percent (4.7 percent) of the NLSY infants were both low birthweight and preterm. NLSY data consistently confirmed those factors well known to influence pregnancy outcome.***

Prominent within the high stress group are jobs of stitchers and sewers in the apparel industry, waiters and food counter workers, nursing aides and orderlies, and cashiers, especially in grocery stores. Low demand, high control work included accountants, mail order salesmen, and bookkeepers. Low demand, low control workers included typists, a variety of cleaning personnel, and household child care workers. High demand, high control occupations included nurses, receptionists, and restaurant managers.†

Impact of Job Characteristics

Women working during pregnancy in jobs characterized by high psychologic demands and low job control were two times as likely to deliver a low birthweight, preterm infant as women working in low exertion jobs (5.1 percent vs 2.6 percent). Similarly, the frequency of low birthweight infants was 1.67 times higher among high stress workers and the mean birthweight of their infants was lower by 78 grams.

Women in high stress jobs were less well educated, had lower prepregnancy weight, and higher smoking frequency***; moreover, almost three-fourths of them in psychologically stressful jobs were also in jobs with high amounts of physical exertion, and physical exertion itself was strongly associated with increased frequency of preterm, low birthweight as well as with the other outcomes (Table 2).

After taking into account the potential confounding effect of job related exertion and other maternal characteristics, we found that the association of job related psychologic stress with preterm, low birthweight, or with any of the other pregnancy outcomes for the sample of women viewed as a whole was no longer important (Table 3).

For those 70 women who stated they did not want to working outside the home after age 35, however, stressful work significantly worsened pregnancy outcomes.

On average their birthweights were almost 500 grams less than infants born to women in less stressful occupations. These differences remained significant when exertion and other maternal factors were considered (Table 4).

Discussion

We found in this study that young women working in jobs characterized by high levels of psychologic demand with little control over the pace and style of response to those demands were somewhat more likely to deliver a preterm, low birthweight infant than women in other jobs, if they were not motivated to continue working.

We considered practically all those factors outside the work experience that are thought to influence pregnancy outcome, and by using an occupational title-based system of job characteristic description, we avoided the possibility of recall bias in determining work stress. The construct of stressful work as being comprised of work with high demands and low control is intellectually appealing; moreover, it, together with the specific instrument used in this study, has been supported in the field of cardiovascular research.[9,16,17]

Despite these strengths, this study does present potential drawbacks in a number of areas. These areas include the nature of the study population, the characteristics of the measures of occupational stress and motivation, and the source of outcome information.

The population used for this study is not representative of all American mothers, but rather represents only those women bearing children during the earlier portion of their potential reproductive years. Such a sample necessarily includes more socioeconomically disadvantaged, unmarried, and minority women than a sample drawn from the entire maternal population. Compared with the women included in the National Family Growth Survey, NLSY mothers are more likely to be Black (21.9 vs 14.1 percent), to have dropped out of high school (29 vs 24.5 percent), and to be unmarried (22.1 vs 17.2 percent).[18] The youthfulness of the sample likely accounts for the somewhat increased preva-

TABLE 2—Pregnancy Outcomes for Women in Jobs of Differing Levels of Exertion (by quartile)*

Job Exertion Level	N	%Preterm/LBW	%LBW	Birthweight (mean in grams)
Quartile 4 (High Exertion)	n = 196	5.07	9.05	3281
Quartile 3	n = 180	7.01	8.83	3232
Quartile 2	n = 232	1.25	4.35	3423
Quartile 1 (Low Exertion)	n = 164	1.03	3.04	3440

*Results not adjusted for differences between workers at different levels of exertion. Singleton births only.

***Data available on request to the author

†A more complete listing of jobs is available on request to the author

TABLE 3—Impact of Stressful Work on Pregnancy Outcomes*

Outcome	Crude Estimate of Effect	Adjusted Estimate of Effect	95% Confidence Interval
Preterm/LBW	Relative Risk = 2.0	Relative Risk = 1.3	.6,3.1
LBW	Relative Risk = 1.7	Relative Risk = 1.4	.75,6.8
Birthweight	Difference = −78 grams	Difference = .1 grams	−81,+81 grams

*Singleton births only; adjusted estimates based on regression models without interaction.

TABLE 4—Impact of Stressful Work on Pregnancy Outcomes, Accounting for Maternal Motivation to Work

Outcome	Crude Frequency of outcome	Adjusted Estimate of Effect (95% CI)
Preterm/Low Birthweight		
Does not aspire to work outside home (n = 70)	Low Stress 0% High Stress 8.3%	Relative Risk = 8.4 (1.4,50.2)
Does aspire to work outside home (n = 716)	Low Stress 2.8% High Stress 3.85%	Relative Risk = .64 (.23,1.7)
Low Birthweight		
Does not aspire to work outside home	Low Stress 0% High Stress 18.7%	Relative Risk = 8.0 (1.3,37)
Does aspire to work outside home	Low Stress 5.7% High Stress 6.3%	Relative Risk = .86 (.39,1.89)
Birthweight		
Does not aspire to work outside home	Low Stress 3561 grams High Stress 3062 grams	mean effect = −295 grams
Does aspire to work outside home	Low Stress 3300 grams High Stress 3262 grams	mean effect = +38 grams

lence of low birthweight relative to concurrent national figures of 6.6–7.6 percent.[2] These demographic characteristics may also explain the decreased frequency of maternal work during pregnancy (66 percent of married primigravidas in the sample worked during pregnancy, compared with contemporary national figures of 85 percent).[1] Nonetheless, while not reflecting the entire age spectrum, the NLSY's emphasis on this younger child-bearing population is important in its own right. Almost half (47 percent) of all births occur to women less than 25 years old.[18]

Concerning the measure of occupational stress, problems may exist relating to both the precision of the measure and its validity among this young population. The use of an occupational title-based system leaves open the possibility that the job experience of the particular women in this study may not be accurately reflected in the job characteristics measure. The indicator of job demand is particularly suspect on this score. Although these measures are unlikely to be biased in any particular direction, use of any imprecise system makes identification of an association difficult.[19]

Moreover, the construct of high demand/low control work as constituting work-related psychologic stress may not fully encompass factors resulting in work stress for young women at early stages of their work careers. Other factors, such as salary, shift worked, flexibility in hours, or availability of maternity and sick leave, may play an equally important role for this population.

The measure of maternal motivation toward work used in this study is not intended to validly predict a woman's employment status at age 35; rather, this measure is used to indicate the woman's current attachment to working. This measure, unfortunately, is not stable even over short periods. Although 9 percent of women in 1979 and 10 percent in 1982 were classified as not intending to remain in paid work, the correlation between these two groups is low (r = .15, kappa = .15).

Finally, all of the outcome measures in this study are based on maternal recall. Although this defect is not found in other studies relating to work and pregnancy outcome, we found the maternal responses very plausible. Mean White and Black birthweights in the NLSY sample, for example, are extremely close to the estimates from the National Natality Survey (3374 grams and 3095 grams for White and Black infants, respectively, in the NLSY, compared with 3391 grams and 3099 grams for White and Black infants in the NNS).[20] Moreover, these data support most of the well known associations with birthweight, low birthweights, and prematurity, such as race, smoking, maternal weight gain, and prior pregnancy experience.

Assuming that our findings are not simply accounted for by these methodologic limitations, why should the impact of high demand/low control work be so strong among women not intending to remain in the paid work force? We can speculate that such women may feel compelled to work, resent it, and that a demanding and restrictive job provides the "last straw." Conversely the over arching desire to remain a part of the gainfully employed population may limit the influence of high demand/low control work on the majority of women who are motivated to remain workers.

Maternal work during pregnancy is likely to be a balance between helpful aspects, such as financial remuneration and social contact, and harmful components. This study sought to determine whether work-related psychosocial stress is one of these harmful aspects. We found that for those relatively few women not desiring to remain in the paid work force, jobs characterized by high demand/low control were independently associated with adverse pregnancy outcome. This finding, based on small numbers, must be considered preliminary pending confirmation in larger studies.

For the remainder of working women, high demand/low control work did not seem to influence the outcomes examined. Due to the limitations in the measures used, however, this study does not preclude an important role for work-related psychologic factors in influencing the outcome of pregnancy for these women. Further research is necessary to clarify which aspects of work are stressful for young women, how these interact with the woman's motivation to work, and how these factors together may influence pregnancy outcome.

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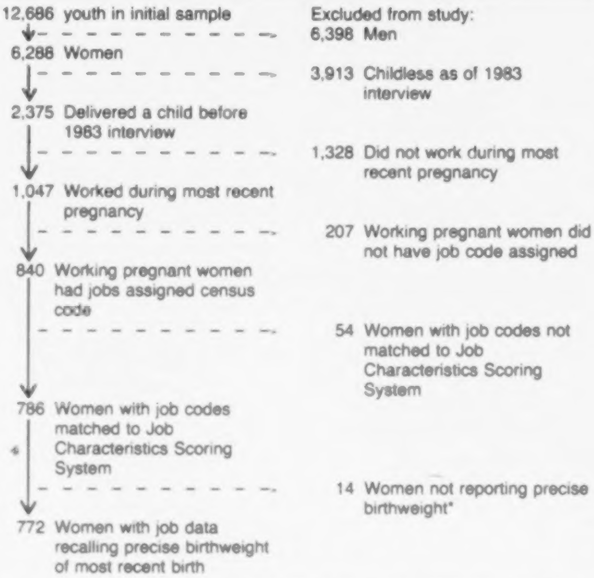
APPENDIX

Derivation of Study Sample from National Longitudinal Survey of Youth


12,686 youth in initial sample
Excluded from study:
6,398 Men
6,288 Women
3,913 Childless as of 1983 interview
2,375 Delivered a child before 1983 interview
1,328 Did not work during most recent pregnancy
1,047 Worked during most recent pregnancy
207 Working pregnant women did not have job code assigned
840 Working pregnant women had jobs assigned census code
54 Women with job codes not matched to Job Characteristics Scoring System
786 Women with job codes matched to Job Characteristics Scoring System
14 Women not reporting precise birthweight*
772 Women with job data recalling precise birthweight of most recent birth

*12 of these women do report whether their infant weighed more or less than 5.5 pounds.

Qualitative Health Research Conference: Call for Abstracts

An international, interdisciplinary conference is being planned for next year to explore issues and developments in qualitative methods and to examine the latest qualitative health research. The conference will be held February 22–23, 1991 at West Edmonton Mall, Edmonton, Alberta, Canada.

Interested researchers are invited to submit abstracts, not to exceed 150 words, for consideration. Send three copies of the abstract—including author(s) name(s), address, institutional affiliation, telephone and fax numbers—by April 19, 1990 to: Janice M. Morse, RN, PhD, Professor and National Health Research Scholar, Faculty of Nursing, CSB 3-120, University of Alberta, Edmonton, Alberta, Canada T6G 2G3. Dr. Morse is also the contact for further information.

Effects of Workplace Smoking Bans on Cigarette Consumption

RON BORLAND, PHD, SIMON CHAPMAN, PHD, NEVILLE OWEN, PHD, AND DAVID HILL, PHD

Abstract: A sample of staff working in the Australian Public Service (n = 2113) were surveyed two to four weeks before a mandated total ban on workplace smoking was introduced, and again five to six months later. Among the 391 smokers on whom complete data were available, the workplace smoking bans were associated with reduced rates of smoking, particularly among heavier smokers where the reduction in consumption was over 25 percent. (*Am J Public Health* 1990; 80:178–180.)

Introduction

The evidence for effects of workplace bans on smoking cessation is mixed[1–5]; the reduction in prevalence may be no greater than the spontaneous rate of smoking cessation. Three studies which investigated consumption in continuing smokers, however, found evidence of reductions.[1,3,5]

We examined the impact on smoking behavior of the introduction of a workplace smoking ban in the Australian Public Service, Australia's largest employer group, with over 170,000 employees. The introduction of the ban was announced over a year before its implementation deadline and was followed by publicity, pre-deadline phasing in of restrictions at some worksites, and availability of programs to help smokers stop smoking or to control their consumption. This paper reports evidence of changes in smoking behavior in a longitudinal sample of smokers from work settings with varying levels of restrictions on smoking prior to the mandated ban.

Method

A total of 2,113 employees of the Australian Public Service (from six Departments spread across three cities and 44 locations) was surveyed two to four weeks before the deadline for implementation of the ban, and again five to six months after its introduction. There were 4,215 respondents to the initial survey (a 79 percent return rate). However, 26.3 percent did not include their names on the initial questionnaire and another 11.5 percent had left the public service or were on extended leave, leaving 2,623 whom we attempted to recontact. A total of 2,169 matched surveys were returned (a return rate of 83 percent). However, 56 of these were not usable, leaving the final sample of 2,113.

Preliminary analyses revealed that those who completed both surveys were less likely at the pre-test to have been smokers (22.7 percent) compared to 27.6 percent for those who completed the initial survey only. There were slightly more males in the final sample (55 percent) compared to the initial sample (51 percent), but there was no difference in mean age or in overall daily cigarette consumption.

To explore the effects of the ban on workday cigarette consumption, only the 391 smokers who reported being smokers at both surveys were considered. Excluded from this group were 39 occasional smokers who reported in the initial survey that they did not smoke on workdays and five who did not provide consumption estimates.

The initial questionnaire, completed two to four weeks before the deadline for the introduction of the ban, dealt with smoking status, the extent of current restrictions in their workplace, and the manner in which any existing restrictions were brought about. Smokers were asked to estimate the number of cigarettes they usually smoked on both work and non-work days, and to recall the number of cigarettes they smoked in the previous 24 hours, divided into seven time periods (before work, while working indoors, while working outdoors, during coffee breaks, at lunch time, in the hour immediately after work, and in the rest of the evening). They were also asked about their attitudes to smoking and to smoking bans, but these data are reported elsewhere.[6] A similar questionnaire was administered at follow-up five to six months later.

Results

Of the matched sample of 2113, 492 (or 23.3 percent) were current smokers at the time of the initial survey and 471 (or 22.3 percent) were smoking at follow-up, a reduction of 21 smokers. Fifty-seven smokers at the time of the initial survey were not smoking at follow-up, and 36 previous non-smokers reported smoking. This 1 percent reduction in prevalence over the six month period is about the estimated yearly reduction in prevalence of smoking for Australians.[7] Of those who stopped, 58 percent had done so more than four months before completing the follow-up questionnaire, i.e., closely following the introduction of the ban.

To study the effects of pre-deadline restrictions on consumption, restrictions were divided into those where the person could smoke at their work station (43 percent of respondents), and those where they should not (57 percent). As can be seen from Table 1, the pre-deadline daily consumption difference between those who reported that smoking at their work station was not permitted and those where it was permitted was 4.5 cigarettes per day. Table 1 also shows that the mean reduction in smoking following the ban was 5.2 cigarettes where there had been no pre-deadline work station restrictions, and 1.9 cigarettes where there were such restrictions. This smaller reduction in the latter group should be seen in the light of their lower pre-deadline consumption.

To further explore the consumption reduction in respondents for whom there had been no restrictions prior to the mandated ban, levels of reported usual workday consumption before the ban were divided into three categories (light, moderate, and heavy) and the change from reported pre-ban workday consumption to post-ban consumption organized by usual pre-ban consumption (Figure 1). While light smokers did not change consumption, moderate smokers reduced by an average of 5.8 cigarettes per day (29.1 percent) and heavy smokers reduced by 7.9 cigarettes (26.6 percent).

Address reprint requests to Ron Borland, PhD, Center for Behavioral Research in Cancer, Anti-Cancer Council of Victoria. 1 Rathdowne Street, Carlton South, Victoria 3053, Australia. Dr. Hill is also at that address; Dr. Chapman is with the Department of Community Medicine, University of Sydney; Dr. Owen is with the Department of Community Medicine, University of Adelaide. This paper, submitted to the Journal March 13, 1989, was revised and accepted for publication June 26, 1989.

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TABLE 1—Mean Daily Cigarette Consumption before and Six Months after the Workplace Smoking Ban was Introduced, According to Restrictions Previously Imposed

Time of Survey	Pre-ban Smoking Restrictions: Banned at Work Station (n = 221)	Not Banned at Work Station (n = 170)	Difference (between subjects effect)
Before ban	16.01	20.54	4.53 (2.45, 6.61)
After ban	14.16	15.36	1.20 (−0.55, 2.95)
Before-after difference (within subjects effect)	1.85 (1.00, 2.70)	5.18 (3.94, 6.42)	

NB: 95% confidence intervals in parentheses.

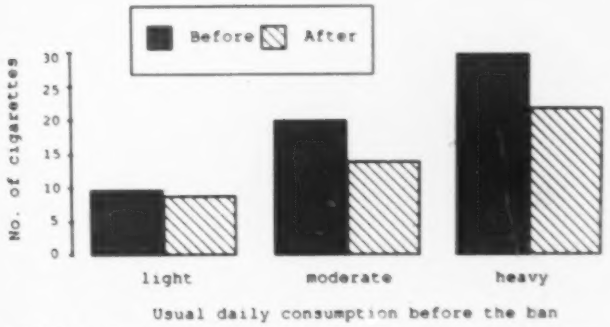

FIGURE 1—Mean Cigarettes Smoked per Day before and after the Ban among Previously Light, Moderate and Heavy Smokers
Light = 1–14 cigarettes/day
Moderate = 15–24 cigarettes/day
Heavy = 25 or more cigarettes/day

As can be seen from Table 2, smokers reduced their average cigarette consumption while working inside by 7.0 cigarettes per day. This was partly compensated for by average increases of 0.7 of a cigarette smoked outside during working hours, and of less than half a cigarette both before work and in the hour after work. The pattern of change was generally consistent for the three smoker categories, except that heavy smokers tended to smoke more outside while working, and the light smokers showed evidence of compensation in the "rest of the evening."

Discussion

Although some smokers may have used the ban's onset and/or the increased availability of smoking cessation resources as stimuli to stop, overall prevalence in the sample changed very little; it is not clear that this resulted in increased rates of smoking cessation compared with normal community rates. The introduction of the workplace smoking ban was associated with reduced levels of cigarette consumption, however. Small increases in rates of smoking outside of the work environment did not compensate for the enforced reduction. Comparison of smoking behavior under two prior levels of restriction on smoking corroborate the evidence that workplace smoking bans are associated with reduced overall levels of smoking, as do data, not reported here,* which showed similar drops in consumption among a small sample who completed 24-hour smoking diaries both before and after the ban.

Relatively little compensatory smoking occurred, and there is no evidence of greater compensation in heavier smokers, who are presumably the most addicted. The fact that smokers did not smoke more at coffee and lunch breaks suggests that the inconvenience of leaving the workplace to smoke was a sufficient obstacle to increased consumption. However, because we did not assess blood nicotine or cotinine levels directly, and did not ask for reports of puff topography, it is not possible for us to show that compensatory changes in puff topography did not occur. Smokers could have increased puff frequency, depth of inhalation, or amount smoked as compensatory mechanisms. However, the environmental restraints imposed by a workplace smoking ban do not favor compensation by changed smoking topography because they reduce opportunities to space cigarettes throughout the day, and compensatory smoking patterns

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*Available on request to author.

TABLE 2—Mean Changes in Cigarette Consumption by Usual Smoking Level Reported in the Initial Survey for Those Who Could Smoke at the Work Stations at that Time, Broken Down by Time of Day

	Pre-testing smoking level: Light n = 48	Moderate n = 62	Heavy n = 60	Total n = 170
Before work	0.17 (−0.11, 0.45)	0.31 (0.07, 0.55)	0.80 (0.38, 1.22)	0.44 (0.25, 0.63)
Coffee breaks	−0.04 (−0.44, 0.36)	−0.16 (−0.60, 0.28)	0.07 (−0.42, 0.56)	0.05 (−0.20, 0.30)
Working inside	−2.83 (−3.60, −2.07)	−6.42 (−7.58, −5.26)	−10.97 (−12.79, −9.15)	−7.01 (−7.93, −6.09)
Working outside	0.46 (0.10, 0.82)	0.24 (−0.33, 0.81)	1.37 (0.50, 2.24)	0.70 (0.32, 1.08)
Lunch	0.38 (0.08, 0.67)	−0.13 (−0.37, 0.11)	0.13 (−0.33, 0.60)	0.11 (−0.09, 0.31)
Hour after work	0.46 (0.15, 0.77)	0.31 (0.01, 0.61)	0.55 (0.21, 0.89)	0.44 (0.62, 0.26)
Rest of evening	0.48 (−0.46, 1.42)	0.05 (−0.84, 0.93)	0.13 (−1.18, 1.44)	0.20 (−0.41, 0.81)
Total	0.94 (−2.51, 0.63)	−5.81 (−7.32, −4.30)	−7.9 (−10.62, −5.21)	−5.18 (−6.41, −3.95)

NB: 95% confidence intervals in parentheses.

would result in increased exposure, primarily in the evening when most cigarettes are already reported to be consumed. Nevertheless, if there were compensatory changes, then inferences about the benefits of the reduced consumption would be weakened.

Heavy smokers appear to benefit most from a workplace ban. As heavy smokers typically find smoking cessation most difficult,[8] workplace smoking bans may provide a means by which they can reduce their consumption to less dangerous levels. Although controlled smoking has not been shown to be a generally sustainable strategy under normal conditions,[9] the imposition of environmental restrictions may make long-term controlled smoking more viable in itself, and a useful way station for those who would eventually like to stop smoking completely.

To estimate the impact the ban may have on overall cigarette consumption, we extrapolated our results to the entire Australian Public Service. We estimate that 24.7 percent of 170,000 public servants or about 42,000 are smokers. Taking the estimate of reduced consumption of 5.2 cigarettes a day from the main sample in Table 1, and assuming 240 work days per year, the resultant reduction in cigarette consumption is about 52 million cigarettes. At the current cost of about 10 cents (Australian) for a cigarette, this represents a loss of $5.2 million a year in retail tobacco sales to staff of the Australian Public Service alone. The public health benefits of these restrictions seem clear, but it is also evident what is motivating the tobacco companies to oppose them.

ACKNOWLEDGMENTS

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NICHD Funds Two Research Centers on Learning Disabilities

The National Institute of Child Health and Human Development (NICHD) has announced the establishment of the first two US research centers to combine the biomedical and behavioral sciences to focus specifically on learning disabilities. One of the new centers, headed by Dr. Bennett Shaywitz, will be at Yale University in New Haven, CT; the other, headed by Dr. Martha Bridge Denckla, will be located at Johns Hopkins University in Baltimore, MD. Support for the two centers totals approximately $1.5 million, with $550,000 going to Yale, and almost $1 million to Johns Hopkins.

The centers were established in response to the recommendations of the Interagency Committee on Learning Disabilities (ICLD), which was established by the Health Research Extension Act of 1985 to review proposals for research programs and facilities related to learning disabilities in children. The US Department of Education (1987) reports that 4.73 percent of all school-aged children (representing almost 1.9 million children) receive special education services for learning disabilities.

The Yale Center will include three projects: attention and conduct disorders in children and their interrelationships; attentional aspects of cognition; and a study of the distribution, types, and stability of reading, arithmetic and attention problems in a representative sample of children selected at age five and followed for two years.

The Johns Hopkins Center proposes a multidisciplinary program; through the convergence of biomedicine and education, investigators hope to discover how and why children develop learning disabilities. The program will include neurologists, psychiatrists, geneticists, statisticians, psychologists, special educators, and communication specialists.

The investigators plan to use molecular genetics technology and magnetic resonance imaging (MRI) to study the origins of learning disorders. The research will focus on the psychological processes related to brain development as well as the effects of educational intervention. To gain a better understanding of Attention-Deficit-Hyperactivity-Disorder, investigators will study the visual, spatial and motor deficits in families with a history of known genetic defects. Methodological approaches will include neuropsychological analysis and MRI correlates of intervention-resistant reading disabilities.

Precision and Accuracy of a Portable Blood Analyzer System during Cholesterol Screening

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SALVATORE GAGLIANO, BS, AND CHARLES E. SPARKS, MD

Abstract: The precision and accuracy of two Kodak Ektachem DT-60 portable blood analyzers were assessed in a model (research) cholesterol screening program in Rochester, New York. Between June and October 1987, a total of 8,573 people underwent a cholesterol screening held in a movable trailer. A wide variety of temperature, humidity, and other potentially adverse conditions were encountered during the screening period. Between-run coefficients of variation ranged from 1.9 percent to 4.8 percent per month; average bias compared to a Reference Laboratory method ranged between +0.2 percent and +2.0 percent. Both precision and accuracy met currently recommended standards for cholesterol testing in the United States. (*Am J Public Health* 1990; 80:181–184.)

Introduction

In 1985, the National Institutes of Health initiated a National Cholesterol Education Program intended to inform health professionals and the general public about the significance of high blood cholesterol levels.[1] It is now recommended that cholesterol testing be performed on all adults in the United States.[2] Interest in blood cholesterol coincides with the recent availability of analytical systems that allow cholesterol concentration to be measured in a matter of minutes with small volumes of whole blood, serum, or plasma. Some of these systems are portable and could facilitate almost-immediate feedback to screening participants. The instruments appear to be highly acceptable to consumers because they avoid venipuncture in favor of fingerstick blood sampling.[3] While found to be reliable when evaluated in physicians' office or laboratory settings, there is little information concerning the portable blood analysis systems over extended time periods under field conditions.[4–6]

A report from the Laboratory Standardization Panel of the National Cholesterol Education Program[7] indicates that well-established clinical laboratories may lack acceptable accuracy in cholesterol measurement. Given the problems that have been noted with major laboratories in cholesterol testing, it is critically important to assess the reliability of the portable testing systems under true field conditions before they are widely recommended for general use in public screenings. We report such an assessment.

Methods

The Rochester, New York Model Cholesterol Screening Program

A model cholesterol screening program funded as a demonstration research project by the National Heart, Lung, and Blood Institute was conducted in the metropolitan Rochester, New York, area during a 13-week period from June to October 1987. Ten grocery stores served as the sites of the cholesterol screenings, and 8,573 screenees participated.

The Analytical System

The Kodak Ektachem DT-60 was evaluated for total cholesterol determinations. Cholesterol measurement on this system utilizes a dry, multilayered, self-contained analysis element coated on a clear polyester support ("slide"). Cholesterol esters are hydrolyzed to cholesterol, which is then oxidized to generate hydrogen peroxide. The hydrogen peroxide oxidizes a leuco dye in a peroxidase-catalyzed reaction to produce a colored compound. Colorimetric measurement by reflectance spectrophotometry provides the basis for the final cholesterol concentration value. A 10 microliter drop of plasma or serum is deposited on the slide by a single-tipped battery-driven pipette and is evenly distributed on the slide by a spreading layer. Once spotted with plasma or serum, the slide is manually advanced into the analyzer for further processing. Assay time is slightly more than five minutes; six samples may be loaded in the analyzer simultaneously, and maximum throughput is about 65 cholesterol slides per hour. Results appear both in a visual display and on a hard copy.

Two DT-60 Analyzers (designated here DT-60A and DT-60B) were purchased from the regular production line. The instruments were operated by a screening team of 20 people who were trained by supervisory personnel. None of the screeners had prior laboratory or chemistry experience.

Training and Prescreening Operation

Screening supervisors learned to operate the instruments during a standard 3 to 4-hour training period conducted by Kodak sales personnel. For each of 10 days prior to the actual screening program, two supervisors prepared the instruments for use and ran four sets of reference materials in pairs, along with actual patient-derived blood samples, twice each day. Thus, each day, each reference sample was run four times on each analyzer. These reference values were compared to target values provided by the manufacturer and were used to establish calibration of each DT-60. Precision data were assessed to determine readiness for actual field operation. The supervisors then trained the screening team, each member receiving 3 to 4 hours of practice and training prior to field operation of the instruments. In the field, the screening team was supervised closely, especially for the first 2 to 3 weeks to avoid technical errors.

Environmental Conditions

Screenings were held in an 8′ × 35′ construction trailer equipped with a room air conditioner. The trailer was not

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heated, and no attempt was made to standardize trailer temperature or humidity. The trailer was allowed to simulate the wide variety of temperature and moisture conditions that might be encountered in community screening programs. On one potential screening day in early October, when outside temperatures reached below 10°C, both DT-60 analyzers failed to reach operating temperature, and screenings had to be canceled. No other obvious malfunctions were noted due to environmental conditions in the construction trailer. The instruments were started and shut down anew each day.

Precision

Instruments were calibrated using Kodak supplies as described by the manufacturer for the DT-60. At the beginning of each day's operation, a series of three separate commercially prepared reference samples was run on both DT-60 Analyzers to assess continued evidence of calibration. Screenee samples were analyzed either on DT-60A or DT-60B, whichever was ready first. Samples were run in duplicate when the first determination for total cholesterol was <3.35 mmol/L or >7.25 mmol/L. Reference material values were "rejected" when a reference value was greater than two standard deviations from the target value, as determined during the prescreening period on replicates of reference materials. The "rejected" specimen was repeated and, if still out of range, the operations protocol specified discontinuation of analyses until the instrument was recalibrated. After analysis of 10 screening samples on a given instrument, a low-range reference sample was run. Following 10 more screening samples, an intermediate-range reference sample was run, and following an additional 10 screening samples, a high-range reference sample was run. Thus, after 30 screenee samples, a series of three reference samples was completed. This procedure was repeated throughout a day's run on both DT-60 instruments. Between-run precision only is reported here.

Accuracy

Accuracy of the DT-60 instruments was assessed using patient samples (venous serum) obtained from hospitalized patients. Fresh samples were split and run on DT-60A, DT-60B, and by the Abell-Kendall Manual Reference Method in a CDC-standardized laboratory. Paired runs were performed three times during the screening program, approximately six weeks apart. Sixty-eight samples were tested on all three instruments/methods. Samples were run on the DT-60s by usual screening team members at the end of the day's screening work.

Results

The precision data during the 10 days of practice (or training) prior to actual screening in the field are shown in Table 1. While coefficients of variation (CV) for the entire 10-day period, for each instrument (DT-60A and DT-60B), were overall quite acceptable, especially in the clinically important values above 5.15 mmol/L, accuracy of the instruments was less acceptable. For example, the CV for the reference sample target of 7.00 mmol/L was 2.22 percent for DT-60A and 2.59 percent for DT-60B over the 10 days. However, mean values on both instruments fell below target values throughout the entire range of reference samples studied. Thus, we considered the instrument/operator team to be acceptable prior to screening in the field in terms of precision, but both instruments were felt to require recalibration before actual field use to correct for the low bias we observed on reference materials during the 10-day prescreening trial runs.

Precision for DT-60A, on a monthly basis during the actual screening program, is reported in Table 2. Coefficients of variation never exceeded 4.8 percent and were typically 3 percent or less. Data for DT-60B were similar and therefore are not repeated. It should be noted that CV values did not substantially improve with either instrument as the experience of the screeners increased.

The results of linear regression analysis comparing each DT-60 separately against the Abell-Kendall Manual Reference Method data are reported in Table 3. As indicated, each DT-60 correlated highly with the Abell-Kendall values (r = 0.99 for each), and both showed a slightly negative proportional bias (linear slope <1.0) but a positive y intercept. Table 4 reports the mean and average bias of each DT-60 compared with the Abell-Kendall Reference results for all patient samples run (total cholesterol by reference method ranges from 2.75 mmol/L to 9.55 mmol/L). Figure 1 is a Bias Plot comparing the DT-60 instruments with the matching Abell-Kendall Reference Results. The plot provides visual confirmation of the bias data reported in Table 4.

Discussion

The major criteria by which to assess reliability of cholesterol testing are precision and accuracy.

Precision is defined as the ability of an analytical system to obtain comparable values on repeated testing, typically

TABLE 1—Precision during 10 Days of Pre-Screening/Training Period

Instrument	Number of Observations	Target Cholesterol Value of Reference Material (mmol/L)	Minimum Cholesterol Value Observed (mmol/L)	Maximum Cholesterol Value Observed (mmol/L)	Mean of Cholesterol Observations (mmol/L)	Coefficient Vairation (CV) (%)
DT-60A	40	2.60	2.25	2.70	2.50	3.87
	40	3.15	2.55	3.15	3.00	4.23
	40	7.00	6.45	7.10	6.90	2.22
	40	7.60	6.70	7.60	7.15	2.59
DT-60B	40	2.60	2.05	2.75	2.55	5.50
	40	3.15	2.80	3.25	3.05	3.13
	40	7.00	6.40	7.20	6.90	2.59
	40	7.60	6.80	7.55	7.20	2.33

To convert mmol/L to mg/dl cholesterol, divide by 0.02586

TABLE 2—Precision in a Field Study of Model Cholesterol Screening (DT-60A)

Month	QC Pools Target Value (mmol/L)	Mean Value (mmol/L)	DT-60A	
			(n)	CV (%)
#1	2.65	2.65	69	2.8
	6.00	6.05	67	2.6
	7.25	7.30	65	4.8
#2	2.70	2.70	100	3.0
	6.30	6.15	82	2.7
	7.70	7.50	79	3.4
#3	2.70	2.70	88	2.8
	6.30	6.25	81	2.8
	7.70	7.60	72	1.9
#4	2.70	2.70	92	3.3
	6.30	6.35	86	2.7
	7.70	7.65	76	2.3

TABLE 3—Linear Regression Equations for DT-60 Results Compared to Abell-Kendall Reference Method Cholesterol Results

Analytical Instrument	Sample Size (n)	Slope	Y-Intercept	r
DT-60A	68	0.952	14.44	0.99
DT-60B	68	0.946	13.20	0.99

TABLE 4—Mean and Average Bias of Instruments versus Abell-Kendall Method for Cholesterol

Analytical Instrument	Sample Size (n)	Mean (mmol/L)	Average Bias Compared to Abell-Kendall Reference Method**
Abell-Kendall* (Reference Method)	68	6.10	—
DT-60A	68	6.20	+1.3%
DT-60B	68	6.10	+0.2%

*range of values 2.75 mmol/L to 9.55 mmol/L

**Biases calculated using cholesterol values originally determined in mg/dl.

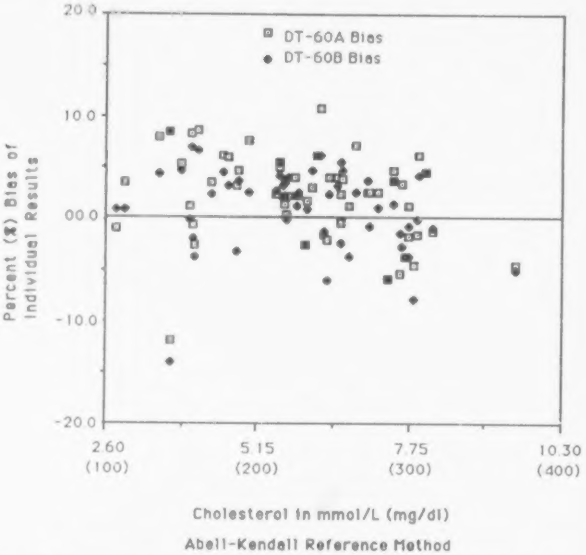

FIGURE 1—Bias Plot Comparing Individual DT-60A and DT-60B Results against the Cholesterol Concentration Determined by the Abell-Kendall Reference Method on Venous Serum Split Samples

expressed by the *coefficient of variation* (CV), defined as one standard deviation divided by the mean of repeated determinations × 100. The *total* (or between-run) coefficient of variation is a measure of a method's variability over many days, whereas *within-run* precision is a measure of variance during a single run. Lippel, *et al*, noted that total precision is usually worse than within-run precision for cholesterol because there are many more sources of variability from run-to-run than would be likely within a single run.[8] The National Cholesterol Education Program has advised that the total CV should be 5 percent or less, and most laboratories now participating in the College of American Pathologists standardization program report total CV under 3.5 percent.[7]

Accuracy is defined as agreement of the test result with the "true value." The most frequently used cholesterol reference method is the manual Abell-Kendall method utilized by the Centers for Disease Control (CDC), generally assessed by comparing cholesterol values obtained by the test method to those of a laboratory whose accuracy has been "traced" back to an accepted reference method, such as that available at the CDC.[7] The Laboratory Standardization Panel of the NCEP has recommended that laboratories achieve an *average* bias less than or equal to ±3 percent within the next five years.[7] Average *bias* is determined by subtracting the mean of a *group* of patient samples run on a new method from the mean of those samples determined using the reference method and then dividing the difference by the mean of the reference method.

Portable cholesterol testing instruments have been available since approximately 1979,[9] and a number of precision and accuracy assessments using portable testing devices have been reported.[4,5,6,10-13] However, previous studies have generally not been performed in screening settings where frequent movement of potentially sensitive equipment, large numbers of different screening personnel with varying degrees of laboratory experience, large patient volume, and changes in environmental temperature and humidity could influence the reliability of test results. The specific purpose of this study was to simulate a community mass screening program and to assess laboratory measurement reliability in a true field setting over an extended time period.

Under the conditions of this study, we noted acceptable results for both precision and accuracy during four months of screening for total cholesterol using either of two DT-60 analyzers we had purchased. Variability remained stable over the entire screening period suggesting that CVs of 3.0 percent may be the best one can achieve under conditions similar to those we encountered. These values for precision and accuracy fall within those recommended as goals for clinical laboratories to achieve by 1993.[7] Thus, assuming operational conditions no worse than those we encountered, one could expect this analysis system to provide both precise and accurate total cholesterol determinations in a public screening setting.

Caution should be exercised in extrapolating results of this study to all screening programs and settings. The results we obtained are probably representative of the range of precision and accuracy values that screening programs might achieve, but a number of variables in other programs could lead to differences. For example, each analyzer is slightly different, as our results showed. Even when instruments are calibrated using identical materials and operated by the same personnel, results from one instrument to another will vary. We would recommend, therefore, that screening personnel perform assessments of instrument accuracy and precision before undertaking public screenings and ensure that these variances are within accepted laboratory standards. We do not recommend that screening teams rely exclusively on reported biases or precision levels since, as the literature now clearly shows, these may vary from instrument-to-instrument and from operator-to-operator.[6,12,14] These results also cannot necessarily be extrapolated to other types of portable blood analyzers such as the Reflotron by Boehringer-Mannheim or the Vision system by Abbott. To date, reports based on the Reflotron have suggested that neither accuracy nor precision are as acceptable in screening as the results reported here.[15,16] No data have been reported concerning the Abbott system in a screening situation.

In conclusion, we found that a trained group of screening personnel using a portable blood analyzer system yielded acceptable precision and accuracy in a demanding screening-type field setting. We expect that similar results could be achieved by other groups when utilizing equally rigorous surveillance of results and comparable instrumentation.

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Report Shows 16 Alzheimer's Medicines in Development

A new report has been released by the Alzheimer's Association and the Pharmaceutical Manufacturers Association (PMA) showing that 16 medicines are being developed by 11 pharmaceutical companies to treat Alzheimer's disease (AD) and its related symptoms.

The report, titled *Alzheimer's Medicines in Development*, lists chemical and trade names of each drug, as well as its developer and development status. All the medicines in the report are in human testing in the United States.

Alzheimer's disease costs Americans close to $80 billion a year and currently affects an estimated 4 million Americans.

Although none of the test medicines is expected to be a cure, they do represent important progress toward easing the symptoms of Alzheimer's disease.

A copy of the report may be obtained by writing the Alzheimer's Association, 70 E. Lake Street, Suite 600, Chicago, IL 60601; or the Pharmaceutical Manufacturers Association, 1100 15th Street, NW, Suite 900, Washington, DC 20005.

Demographic and Dietary Determinants of Constipation in the US Population

Robert S. Sandler, MD, MPH, Matthew C. Jordan, MS, and Brent J. Shelton, BS

Abstract: We investigated the association between self-reported constipation and several demographic and dietary variables in 15,014 men and women 12–74 years of age examined between 1971–75 during the first Health and Nutrition Examination Survey. Overall, 12.8 percent reported constipation. Self-reported constipation correlated poorly with stool frequency. Nine percent of those with daily stools and 30.6 percent of those with four to six stools/week, reported constipation. Constipation was more frequent in Blacks (17.3 percent), women (18.2 percent), and those over age 60 (23.3 percent); after adjusting for age, sex, and race it was more prevalent in those with daily inactivity, little leisure exercise, low income, and poor education. Constipated subjects reported lower consumption of cheese, dry beans and peas, milk, meat and poultry, beverages (sweetened, carbonated and noncarbonated), and fruits and vegetables. They reported higher consumption of coffee or tea. They consumed fewer total calories even after controlling for body mass and exercise. (*Am J Public Health* 1990; 80:185–189.)

Introduction

Constipation is a significant problem in the United States with over 330 million dollars spent each year on over-the-counter laxatives.[1] Our understanding of both the epidemiology and pathophysiology is limited.

Studies of constipation are handicapped by lack of a standard definition of the condition. Connell, *et al*,[2] inquired about the bowel habits of 1,055 persons not seeking medical advice in an industrial community. They found that 99.4 percent had three or more bowel movements per week. Based on this work, individuals with fewer than three stools per week are usually considered to be constipated. While physicians and investigators often define constipation by stool frequencies that lie below the usual range, healthy subjects tend to define constipation in terms of function (straining) and stool consistency (hard stools).[3]

Because of the lack of a standard definition, because bowel habits change with time, and because constipation depends on age and sex, the range of reported prevalence is quite broad. In a group of healthy young adults not seeking health care, Sandler and Drossman found that 7.3 percent reported that they were constipated more than 25 percent of the time.[3] When constipation was defined as straining more than 25 percent of the time, Thompson and Heaton found that 10.3 percent of subjects were constipated.[4] In a prospective study of 1,064,004 men and women who were surveyed by American Cancer Society volunteers, 18.5 percent of men and 33.7 percent of women reported "constipation."[5]

Sonnenberg and Koch have recently performed detailed analyses of constipation using data from the National Hospital Discharge Survey, the National Health Interview Survey, the National Ambulatory Medical Care Survey, and the Vital Statistics of the United States.[6] Constipation was the most frequent digestive compliant in the United States with an estimated four million people complaining of "frequent constipation." There were over 92,000 annual hospitalizations with constipation listed among the hospital discharge diagnoses, and 900 deaths from diseases associated with or related to constipation. Using data from the National Disease and Therapeutic Index, the same authors reported 2.5 million physician visits for constipation.[7] The data are limited in several ways. Most people with constipation do not visit physicians, are not hospitalized, and do not die. The reliability of the prevalence estimates for chronic constipation based on the National Health Interview Survey are uncertain because one member of the household provided health information for all members of that household. The published reports of constipation have not attempted to link constipation with dietary and lifestyle factors that might be responsible for this symptom.

The first National Health and Nutrition Examination Survey (NHANES-I) was conducted by the National Center for Health Statistics between 1971–1975.[8] The study involved a nationwide probability sample of 20,749 persons. In addition to detailed socio-demographic information, the survey included data on dietary intake, exercise, and bowel dysfunction. We took advantage of unique data available from the NHANES-I to describe the distribution and determinants of constipation in the United States.

Methods

The NHANES-I was a large-scale, multistage probability sample designed to closely reflect the US non-institutionalized civilian population, 1–74 years of age. The design and methodology have been described in detail elsewhere.[6] The analyses reported here involve 15,014 Black and White persons older than age 12 at the time of the survey, who completed both an examination and a dietary interview.

Constipation, in the present study, was based on responses to the question: "Do you have trouble with your bowels that makes you constipated or gives you diarrhea?" Those who answered "yes—constipated," were defined as having self-reported constipation and were compared to those who responded negatively to the question. Those with diarrhea (621), constipation and diarrhea (73), and those with missing information (72) were excluded. Subjects were also asked: "How often do you usually have a bowel movement?" Stool frequency was coded by the National Center for Health Statistics as less than two stools per week, two to three per week, four to six per week, one per day, and two or more per day. It was not possible to analyze stool frequencies apart from these categories. An alternative definition of constipation was fewer than four stools per

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week. No time period for self-reported constipation or stool frequency was specified.

Height and body weight were obtained through standardized procedures.[6] Body mass index (BMI) was calculated using the formula BMI = weight (kg.)/height $(m)^2$ to provide an index of weight controlled for height. Daily calories were determined using standardized food composition tables applied to foods in each person's 24-hour dietary recall record. Subjects were also asked the frequency with which they consumed a variety of foods with categories chosen to provide reasonably balanced numbers in each category. Race was dichotomized as White or Black.

Analyses were performed using the Statistical Analysis System (SAS Institute, Cary, NC). The SESUDAAN program, which takes into account the complex sample design of the survey, was used to compute standardized proportions and standard errors.[9] The RTILOGIT program[10] was run under SAS to perform logistic regression analyses that took account of the sample weights and sample design. Age, body mass index, and calories were entered into models as continuous variables. Categorical variables were coded as described in the Tables.

Results

Overall, 12.8 percent (S.E. 0.48) of the United States population reported constipation (Table 1). Self-reported constipation was more common in Blacks (17.3 percent), women (18.2 percent), and those over age 60 (23.3 percent). When we defined constipation as fewer than four stools per week, constipation was again more common in Blacks and women. The elderly did not consistently report fewer stools.

There was poor correlation between the reported stool frequency and self-reported constipation (Table 2). Only 67.8 percent of those with fewer than two stools per week reported constipation. Nine percent of those with daily stools, and 30.6 percent of those with four to six stools per week considered themselves constipated. Of the 180 persons with less than two stools per week, 137 (76 percent) were women, supporting the fact that constipation is more common in women.

Self-reported constipation was most prevalent in those with daily inactivity and little recreational exercise with a distinct gradient from the most active to the least active at all ages (Table 3). Those with low income and education were more likely to be constipated within each age category. The effect of marital status was inconsistent. Reported constipation was slightly more common in those living in the South. When we stratified by race and sex (not shown), all of these findings persisted. They also persisted when we compared those with fewer than four stools per week (constipated group) to those with more than four.

TABLE 1—Proportion of Persons with Self-Reported Constipation in the US Population

Group	No.	% Constipated*	Standard Error
Overall	15,014	12.8	0.48
Race			
White	12,178	12.2	0.48
Black	2,836	17.3	1.40
Sex			
Men	6,042	7.0	0.50
Women	8,972	18.2	0.70
Age (years)			
12–29	5,660	8.1	0.54
30–59	5,470	13.9	0.61
60–75	3,884	23.3	1.21

*Percents are US population estimates derived by applying sampling weights to survey data.

TABLE 2—Comparison of Stool Frequency to Self-Reported Constipation

Stool Frequency[a]	No. (%)	% Constipated[b]
<2/week	180 (1.0)	67.8
2–3/week	1,342 (7.6)	58.3
4–6/week	1,884 (10.6)	30.6
1/day	12,442 (70.0)	9.0
≥2/day	1,930 (10.9)	7.0

[a] "How often do you have a bowel movement?"

[b] "Do you have trouble with your bowels which makes you constipated?"

TABLE 3—Proportion of Persons with Self-Reported Constipation in the US Population by Age and Personal Characteristics

	Age (years)		
	12–29	30–59	60–75
	% SEM[a]	% SEM[a]	% SEM[a]
Nonrecreational Exercise[b]			
Very active	6.9 (0.6)	11.7 (0.8)	17.9 (1.2)
Moderately active	8.7 (0.7)	15.9 (0.9)	25.8 (1.7)
Inactive	15.9 (3.1)	17.2 (2.3)	29.0 (3.2)
Recreational Exercise[c]			
Much exercise	6.0 (0.6)	8.9 (1.0)	18.0 (2.4)
Moderate exercise	7.5 (0.8)	12.7 (0.9)	19.3 (1.7)
Little exercise	14.3 (1.3)	17.1 (1.1)	27.8 (1.7)
Income			
$<6,999	9.6 (1.2)	20.6 (1.9)	28.8 (1.6)
7,000–9,999	8.3 (0.8)	14.9 (1.3)	18.4 (2.1)
10,000–14,999	7.9 (0.9)	11.5 (1.0)	17.1 (3.5)
>14,999	6.4 (0.9)	9.8 (1.2)	10.3 (1.9)
Education (years)			
0–6	9.8 (2.6)	25.6 (3.1)	33.5 (2.8)
7–11	7.0 (0.7)	15.8 (1.2)	26.6 (1.7)
12	8.6 (0.8)	12.7 (8.9)	16.6 (2.5)
13+	8.9 (0.9)	10.7 (0.9)	15.4 (2.1)
Marital Status			
Never Married	6.6 (0.6)	20.0 (2.3)	23.4 (3.7)
Married	10.8 (0.8)	12.8 (0.6)	21.8 (1.5)
Region			
Northeast	7.8 (0.8)	13.3 (1.1)	18.4 (2.1)
Midwest	7.1 (0.9)	12.7 (0.9)	23.5 (1.9)
South	9.6 (1.6)	16.6 (1.7)	28.4 (3.2)
West	8.1 (0.9)	13.4 (1.2)	23.0 (2.6)

[a] Standard error of the mean

[b] Subjects were asked: "In your usual day, aside from recreation, how active are you?"

[c] Subjects were asked: "Do you get much exercise in things you do for recreation?"

Table 4 presents data on a number of specific dietary constituents. Constipated individuals reported lower consumption of cheese and cheese dishes, dry beans and peas, milk, fruits and vegetables, meat and poultry, and sweetened, carbonated and non-carbonated beverages, and higher consumption of coffee and tea. There was no consistent relationship with cereal or bread. These associations largely persisted for White men and women and when we defined constipation as fewer than four stools per week. The results for Blacks were not as consistent but were based on much smaller numbers within individual dietary categories.

In order to assess the independent effects of these variables, we examined them in a logistic regression model that also contained terms for age, race, sex, body mass index, and total daily calories. Each individual dietary item was

TABLE 4—Proportion of Persons with Self-Reported Constipation in the US Population by Dietary Constituents Consumed

	Frequency of Consumption[a]				
	low 1	2	3	4	high 5
	% SEM[b]	% SEM	% SEM	% SEM	% SEM
Cheese	15.9 (1.3)	14.1 (1.1)	12.9 (0.6)	11.2 (0.6)	12.6 (1.1)
Cereal	11.9 (0.8)	13.9 (0.9)	11.8 (0.6)	16.1 (1.3)	11.9 (9.4)
Beans	20.6 (1.4)	12.7 (0.7)	12.6 (0.6)	10.7 (0.8)	9.2 (1.0)
Bread	15.4 (6.0)	13.4 (3.9)	13.3 (1.0)	12.6 (0.6)	12.9 (0.6)
Milk	15.3 (1.2)	14.8 (1.7)	14.2 (0.8)	12.4 (0.6)	9.9 (0.8)
Fruit/vegetables	2.6 (2.1)	18.4 (4.1)	14.0 (1.3)	12.3 (0.6)	13.1 (0.7)
Meat/poultry	28.2 (11.1)	22.1 (6.2)	15.9 (1.3)	12.4 (0.5)	5.9 (1.3)
Coffee/tea	7.9 (0.8)	10.3 (1.7)	9.6 (1.1)	13.8 (0.6)	
Beverages	18.0 (1.0)	15.8 (1.0)	11.5 (0.7)	10.2 (0.7)	

[a] Cheese refers to cheese and cheese dishes; beans to dry beans and peas; beverages to sweetened, carbonated and non-carbonated. The categories for the following specific items were: *milk*: (1) never, (2) less than once a week, (3) 1–6 times a week, (4) once a day, (5) more than once a day; *meat and poultry, fruits and vegetables, bread*: (1) never, (2) less than once a week, (3) 1–6 times a week, (4) 1–2 times a day, (5) more than twice a day; *cheese, beans*: (1) never, (2) less than once a week, (3) 1–2 times a week, (4) 3–6 times a week, (5) once or more a day; *cereal*: (1) never, (2) less than once a week, (3) 1–6 times a week, (4) one or more times a day. *Beverages* (sweetened, carbonated and non-carbonated) and *coffee or tea* were ranked in four categories by the NCHS (1) never, (2) less than once a week, (3) 1–6 times a week, (4) once or more a day.

[b] SEM = Standard error of the mean.

statistically significant ($p < 0.001$) in the same direction as the univariate analyses shown in Table 4.

Body mass index was similar for those who reported constipation and those who did not at every age and in both sexes, whereas the constipated group consumed fewer calories (Table 5). In order to assess the independent effects of various demographic and lifestyle factors on reported constipation we performed a logistic regression analysis that simultaneously controlled for all variables in the model and took into account the sample weights and sample design (Table 6). The most important determinant of constipation was sex. Women were 2.4 times more likely to report constipation after controlling for other factors. The other results were in the same direction as the univariate analyses. Constipation was more common in Blacks, the elderly, and those with less education, low income, and inactivity. Marital status and body mass index were not important. Constipated subjects consumed fewer calories even after controlling for the other variables in the model.

TABLE 5—Age-Specific Mean Body Mass Index and Daily Calorie Consumption by Constipation Category[a]

	Age (years)		
	12–29	30–59	60–75
Women			
Body Mass Index[b]			
Constipated	22.7 (0.3)	25.6 (0.2)	27.1 (0.3)
Not constipated	22.2 (0.1)	25.5 (0.1)	26.4 (0.2)
Calories			
Constipated	1698.9 (48.3)	1493.8 (31.1)	1229.4 (33.7)
Not constipated	1762.7 (27.5)	1554.4 (18.8)	1350.5 (23.7)
Men			
Body Mass Index[b]			
Constipated	23.9 (0.7)	26.2 (0.4)	25.3 (0.4)
Not constipated	22.8 (0.1)	26.1 (0.1)	25.5 (0.1)
Calories			
Constipated	2655.5 (125.9)	2195.0 (84.4)	1789.8 (83.1)
Not constipated	2847.2 (39.7)	2440.2 (34.4)	1944.1 (38.6)

[a] Numbers in parentheses represent the standard error of the mean.

[b] Body mass index = weight (kg)/height $(m)^2$

Discussion

Our results are representative of the non-instiutionalized civilian US population. The large size of the study (15,014 subjects) permitted us to look at several demographic and dietary characteristics in order to present a more complete description of constipation. The results confirm other reports and indicate that recreational and nonrecreational inactivity is associated with self-reported constipation independent of dietary intake. Exercise is thought to increase propulsive movements in the large intestine[11] although there are not good studies to support the use of exercise in constipated subjects.[12]

The NHANES-I survey was conducted between 1971 and 1975. There has been a modest decrease in the prevalence of "frequent constipation" between 1968 and 1986.[6] We chose the first NHANES-I because subjects were asked about both stool frequency and self-reported constipation. There were only 180 subjects with fewer than two stools per week, too few to permit detailed analyses. We performed all analyses using stool frequency less than four per week (8.6 percent of the population) as well as self-reported constipation, and found no major differences in the results for diet and demographic features with one exception. The elderly did not report consistently fewer stools than the young even though they reported more constipation.

In the absence of an agreed upon definition of constipation, we believe that self-reports of the symptom of constipation may be preferable to stool frequency since infrequent stools may be asymptomatic and simply represent the lower end of the continuum. Self-reported constipation may refer to several different symptoms including infrequent stools, feelings of incomplete evacuation, straining and hard stools[3] which may, in turn, be the consequence of several discrete pathogenetic mechanisms. It is possible that the differences that we observed in constipation prevalence by age, sex, and race are due to differences in definitions used by respondents. The similarity of our age and sex findings to those based on physician diagnosis, however, argues against this interpretation.[7]

There is value to studying self-reported constipation since it is this symptom that distresses patients. Epidemiol-

TABLE 6—Constipation in the US Population, Logistic Regression Analysis[a]

Variables	Beta	S.E.	Odds Ratio
Female sex	0.8803	0.0007	2.41
Black race	0.2849	0.0008	1.33
Age (10)	0.2063	0.0002	1.23
Nonrecreational activity	0.1813	0.0004	1.20
Recreational activity	0.1522	0.0004	1.16
Married	−0.0090	0.0001	0.99
Body mass index (10)	−0.0279	0.0005	0.97
Calories (1,000)	−0.1391	0.0004	0.87
Income	−0.1369	0.0003	0.87
Education	−0.1503	0.0003	0.86

[a] Adjusting simultaneously for all other variables in model using RTILOGIT[10] to take account of sample weights and sample design. For education, income, nonrecreational activity, and recreational activity, betas (and odds ratios) reflect comparisons between individual categories described in Table 3. Age and body mass index were entered as continuous variables with betas calculated for intervals of 10. Total calories were entered as a continuous variable with the beta reflecting an interval of 1000 calories.

ogists have been reluctant to study conditions for which there may not be pathologic, radiologic, or laboratory confirmation. We need to study the epidemiology of symptoms and symptom complexes, however, since they represent a large part of human morbidity and suffering.[13]

Diet, particularly dietary fiber, has often been regarded as an important determinant of constipation,[14,15] although personality factors may be important as well.[16] The large sample size and detailed dietary data available in the NHANES-I permitted us to examine the association between diet and constipation. We found that those who ate more fruits and vegetables had fewer complaints of constipation. In addition, however, we found that those who ate more of a number of dietary constituents including cheese, dry beans and peas, milk, and meat and poultry also had less self-reported constipation. As one might expect, these individuals also had a higher caloric intake. When we controlled for caloric intake, however, these dietary components were still significant.

One of the most interesting findings was that constipated subjects ate fewer calories. The constipated group was not leaner, and their lower calorie consumption was not completely explained by decreased physical exercise or differences in other demographic characteristics, although these could have been incompletely controlled in the logistic model (residual confounding). The inverse association between calorie consumption and self-reported constipation may offer clues about etiology. Caloric intake is based on body size, physical activity, metabolic efficiency, and net energy balance (i.e., change in body fat or muscle).[17] The mechanisms of metabolic efficiency, including differences in absorption and thermogenesis, are poorly understood, but individual differences must exist since even under controlled conditions some subjects will gain more weight than others eating similar calories.[18]

One might hypothesize that individuals with constipation are more efficient at absorbing or utilizing food in their small intestines. They eat less, present less to their colons, and have more constipation as a consequence. Those free of constipation, on the other hand, are less efficient. Food that escapes absorption enters the large bowel where fecal flora may produce short chain or volatile fatty acids that protect against constipation.

The absorption of dietary constituents by the human small bowel is known to be incomplete. Kramer, *et al*,[19] found that several different dietary components increased the ileal effluent in seven colectomized subjects. The starch moiety of most staple foods (wheat, corn, oats, potatoes) is incompletely absorbed by normal subjects,[20] and there may be considerable variation among individuals. Stephen, *et al*,[21] found that while some individuals absorbed starch well, others failed to absorb 15–20 percent. It is possible that functional constipation could be due to overly efficient handling of the carbohydrate that normally escapes absorption in the small bowel.[22]

In summary, we found that self-reported constipation was more common in women, Blacks, the poor, the elderly, and those with less education and little physical exercise. Those who ate fewer calories were more constipated even after controlling for exercise and body mass index. In order to reconcile all of the available information, one might view constipation as the end result of both exogenous factors (diet and exercise) and endogenous factors (sex hormones, personality, and metabolic efficiency). Further studies on energy balance and absorption in constipated subjects may be important in helping us to understand and treat this common symptom. Such studies will require large numbers and careful consideration of symptom definition.

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Nation's First Medical School Celebrates Its 225th Anniversary throughout 1990

The University of Pennsylvania School of Medicine—the first medical school established in the United States in 1765—marks its 225th anniversary during 1990, at the same time the University notes its 250th anniversary. A year-long celebration is planned to celebrate this landmark in the history of American medicine. A blend of history, contemporary scholarship, and a look at the future will be linked with festivities to make 1990 a memorable year for alumni, faculty, students, and friends.

The continuing theme throughout the year will be the evolution of medicine in the United States and the critical role the University of Pennsylvania School of Medicine played in the process, accentuated by exhibits, lectures, symposia, and exchanges to highlight the major contributions of the school. Three peak celebration periods are planned for the year—occurring in January, May and October.

- The January program emphasized frontiers of science in medicine, and centered around the dedication of a new clinical research building on the site of the former Philadelphia General Hospital. Louis Sullivan, MD, US Secretary of Health and Human Services, delivered the keynote address which focused on national health policy. Other events in the January program included the 27th annual Women in Medicine dinner, and a scientific symposium entitled "Molecular Mechanisms of Disease" featuring three of the world's foremost medical scientists: Philip Leder, MD, Harvard University; Sydney Brenner, MD, Cambridge, England; and Joshua Lederberg, PhD, Rockefeller University.
- The "May Festival" will take place May 16–20, and will include dedication of the Seymour Gray Foundation for Molecular Medicine: Seymour Gray Lecture in Molecular Biology. Keynote speaker will be Michael S. Brown, MD, Nobel Laureate and Professor of Molecular Genetics, University of Texas-Southwestern Medical School. "Medical Alumni Weekend" a major annual event—will be incorporated into the festivities, with numerous social activities, and a scientific symposium entitled "Recent Developments in Cancer Research."

Also in May, continuing medical education credit will be given for a series of faculty-alumni exchanges that are part of the Medical School's 225th anniversary.

- The "Future of Medicine and Health Care" will be the focus of the October festivities, with a symposium on October 12 featuring three keynote speakers: Arnold Relman, MD, Editor, *New England Journal of Medicine*; Samuel O. Thier, MD, President, Institute of Medicine/National Academy of Sciences; and Uwe E. Reinhardt, PhD, Professor of Political Economy, Princeton University.

For further information, contact: Office of the 225th Anniversary Celebration, University of Pennsylvania School of Medicine, 196 John Morgan Building, 36th and Hamilton Walk, Philadelphia, PA 19104; or the Office of Public Information, University of Pennsylvania Medical Center, 2 Blockley Hall, 3400 Spruce Street, Philadelphia, PA 19104-6021. Tel: (215) 662-2560.

Epidemiologic Assessment of Screening Tests for Antibody to Human T Lymphotropic Virus Type I (HTLV-I)

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AND JONATHAN E. KAPLAN, MD

Abstract: We tested 196 sera from a human T lymphotropic virus type I (HTLV-I) risk group (prostitute women) with two commercial "research" enzyme-linked immunoabsorbent assays (EIA) for HTLV-I antibodies. All tested sera were characterized by HTLV-I Western immunoblots and by HTLV-I radioimmunoprecipitation assays. The estimated sensitivities of the EIA tests were 93.8 percent and 100 percent, and the specificities were 98.8 percent and 95.8 percent, respectively, using recommended criteria for seropositivity (requiring reactivity to both *gag* p24 and *env* gp46 or gp61/68). Calculated negative predictive values remained excellent (>99.9 percent and 100 percent, respectively) at lower seroprevalence rates but the positive predictive values were only 7.3 percent and 2.3 percent when calculated for a seroprevalence rate of 0.1 percent. These results emphasize the importance and need for additional HTLV-I supplementary serologic testing when screening populations with low HTLV-I seroprevalence rates. (*Am J Public Health* 1990; 80:190–192.)

Introduction

Human T lymphotropic virus type I (HTLV-I), the first human retrovirus discovered,[1] is causally linked to adult T-cell leukemia/lymphoma (ATL)[2] and has been more recently associated with a progressive myelopathy known as tropical spastic paraparesis (TSP) in the Caribbean and as HTLV-I-associated myelopathy (HAM) in Japan.[3,4] HTLV-I infection is endemic in southwestern Japan, the Caribbean, and in part of Central America and Africa.[5] In the United States, elevated rates of infection have been reported in intravenous drug users[6,7] and in recipients of multiple blood transfusions,[8] raising concerns about the safety of the nation's blood supply. A recent study of random blood donors in eight US cities showed a seroprevalence of 0.025%.[9]

Screening tests for antibodies to HTLV-I were licensed by the Food and Drug Administration (FDA) in November 1988, and the FDA has recommended that these tests be used to screen blood donated for transfusion.[10] The purpose of our study was to evaluate two commercially available HTLV-I enzyme-linked immunoabsorbent assays (EIA)* using a set of sera from a population at high risk for HTLV-I infection, and to estimate the predictive values expected with these tests when testing populations or when screening blood donors from populations with low rates of infection.

Methods

Serum Specimens

The 200 serum specimens used in this study were systematically chosen (every 7th specimen until a total of 200 were obtained) from a set of 1,541 sera from US prostitutes with a high HTLV-I seroprevalence rate (7 percent).[11] All specimens had been frozen and stored at the Centers for Disease Control (CDC) at −20°C and were thawed for this study.

HTLV-I Antibody Testing

All serum specimens were tested by two HTLV-I EIAs (Dupont Co., Wilmington, DE, and Cellular Products Inc., Buffalo, NY); both used viral lysates from HuT-102-infected cells and differed mainly in their respective manufacturer's recommended cutoffs: 0.6 × mean of positive controls for Dupont's and 0.25 × mean of positive controls for the Cellular Products'. Initially reactive sera were retested in duplicate and classified as reactive or nonreactive for the test based on the manufacturer's recommendation. Because we needed to know whether each specimen was positive, indeterminate, or negative (see following), 196 of the 200 serum specimens, including all those reactive in either screening test, were further tested by Western immunoblot (WIB) and by radioimmunoprecipitation assay (RIPA); four specimens, negative on both EIAs, were too small for us to do WIB or RIPA and therefore were excluded from further analysis.

WIB used an HTLV-I-infected cell lysate purified from MT-2 cells (Hillcrest Biologicals, Cypress, Calif).[12] HTLV-I antigen was suspended in a sample buffer, heated at 95–100°C for four minutes, and subjected to electrophoresis in a single well of a 10 percent polyacrylamide gel (PAGE); PAGE-resolved proteins were electrophoretically transferred to nitrocellulose sheets and washed sheets were cut into 3.0 mm strips. Individual strips were incubated with 1:100 dilutions of serum overnight at 4°C, washed, incubated for 60 minutes at room temperature with 5 μg/ml biotinylated goat anti-human IgG (Vector Laboratory, Burlingame, CA). Strips were washed again and then incubated with avidin-biotin-peroxidase complex according to the manufacturer's instructions (Vector Laboratory, Burlingame, CA). Color reactions were developed with 3, 3′diaminobenzidine, nickel chloride and H_2O_2 (Sigma Chemical Co., St. Louis, MO).

For the RIPA, HTLV-I-infected MT-2 cells were metabolically labeled with [^{35}S]cysteine and [^{35}S]methionine (200 uCi of each nucleotide/ml/1 × 10^7 cells), disrupted in lysing buffer, centrifuged, and incubated with the sera. Immunocomplexes were precipitated with protein A-sepharose (Sigma Chemical Co., St. Louis, MO). Samples were analyzed by PAGE followed by autoradiography of dried gels.

In each test run, banding patterns were compared with those of a known positive serum, and a negative serum was included as a negative control. Specificity of banding patterns

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*At the time of this study these tests, which had not been licensed, were approved for research only.

was assessed by comparison with banding patterns on strips that had been incubated with monoclonal antibodies against HTLV-I-specific proteins (p19, p24 for WIB and gp61/68 for RIPA).[13,14]

Definitions

In this study, a seropositive serum sample was defined by the presence of antibodies to *gag* (p24) and *env* (gp46 or precursor gp61/68) on WIB or RIPA.[10] A serum sample with immunoreactivity to at least one HTLV-I gene product but not satisfying these criteria was called indeterminate. All remaining specimens were designated negative.

Calculation of Predictive Values

Negative and positive predictive values were calculated using sensitivity, specificity, and prevalence as outlined in Fleiss.[15]

Results

Of the 196 sera included in the analysis, 16 (8 percent) were HTLV-I seropositive by WIB or RIPA. Of those, seven fulfilled the seropositivity criteria on WIB alone; nine showed only *gag* reactivity on WIB and required RIPA to demonstrate antibody to *env*. Fourteen specimens (7 percent) were indeterminate after WIB and RIPA; all showed reactivity to *gag* gene products only, and no specimen showed reactivity on RIPA that did not also show reactivity on WIB. The remaining 166 specimens were negative.

Of the 16 seropositive specimens, 15 were repeatably reactive by Dupont EIA, for a test sensitivity of 93.8 percent, and all 16 were reactive by Cellular Products' EIA, for a test sensitivity of 100 percent (Table 1). Excluding the 14 serum specimens with indeterminate HTLV-I serologic results, the specificities of the tests were 98.8 percent and 95.8 percent, respectively.

Table 2 shows the positive and negative predictive values of these tests, calculated for different seroprevalence rates; the calculations were based on the sensitivities and specificities estimated in this study: the predictive value of a negative test result increased with lower seroprevalence levels and was >99.9 percent for Dupont's EIA at seroprevalence levels of ≤0.1 percent; the negative predictive value of Cellular Products' EIA was 100 percent for all seroprevalence rates. In contrast, the predictive value of a positive test result decreased with lower seroprevalence rates and was <8 percent for Dupont's and <3 percent for Cellular Products' at seroprevalence rates lower than 0.1 percent.

TABLE 1—HTLV-I Antibody Test Results of 196 Serum Specimens by Repeatable Reactivity in Two HTLV-I Enzyme-Linked Immunoassays

Results of Tests for HTLV-I	No. Repeatably Reactive	
	by Dupont's EIA*	by Cellular Products' EIA
Positive (N = 16)	15	16
Indeterminate (N = 14)	2	10
Negative (N = 166)	2	7

*At the time of the study these tests, which had not been licensed, were approved for research only; the recommended cutoff for the Dupont test has changed with licensure.

TABLE 2—Performance Characteristics of Two EIAs and Resulting Predictive Values for Different Seroprevalence Rates*

	Dupont's EIA		Cellular Products' EIA	
Sensitivity (b)	93.8%		100%	
Specificity (a)	98.8%		95.8%	
if:	PPV[†]	NPV[‡]	PPV	NPV
p = 8%	87.2%	99.5%	67.4%	100%
p = 1%	44.1	99.9%	19.4%	100%
p = 0.1%	7.3%	99.9%	2.3%	100%
p = 0.05%	3.4%	99.9%	1.2%	100%
p = 0.025	1.7%	99.9%	0.6%	100%

*Sera with indeterminate HTLV-I reactivities were excluded from these analyses.

[†] PPV = positive predictive value = $\frac{(b)\,(p)}{(b)\,(p) + (1-a)\,(1-p)}$, where b = sensitivity
a = specificity
p = prevalence

[‡] NPV = negative predictive value = $\frac{a(1-p)}{a(1-p) + (1-b)p}$

Discussion

This paper is not intended to be a comparison of the performance of the two tests. Neither test was licensed by the FDA at the time of the study and the recommended cutoff for the Dupont test has in fact changed with licensure* and is now similar to the Cellular Products' test. Rapid developments in the field may result in additional changes in these tests in the future. Rather, our data are intended to illustrate that both tests have low positive predictive values in groups with low prevalence of infection. Repeating the study with the currently licensed tests would not have changed this conclusion. Our data emphasize the need, with current assays, for more specific supplementary serologic testing before interpreting a sample as seropositive. Further serologic confirmatory testing must precede any decision to counsel a tested individual.

Our study suggests that, in populations with HTLV-I seroprevalence rates <0.1 percent, less than 10 percent of repeatably reactive test results might be found to be truly positive. While quite accurate for populations where sexually transmitted diseases and other infections are prevalent, these estimates of positive predictive value (PPV) must be more cautiously interpreted in screening US blood donors. In the American Red Cross study of 39,898 blood donors, for example, 10 of 68 specimens repeatedly reactive in a screening test were confirmed as positive for HTLV-1 antibody, yielding a PPV of 15 percent.[9] This discrepancy is probably due to a higher specificity of the screening tests in the blood donor population. Specificity could not be calculated accurately in the American Red Cross study because supplementary testing was done only on EIA-reactive specimens, but specificities of 99.3 percent to 99.9 percent (compared with 98.8 percent and 95.8 percent in our study) have been estimated in blood donor populations.[10] Additionally, our lower estimates of PPV could be the result of differences in the cutoff used to interpret these tests, and to technical differences in the supplementary serologic tests. Despite these discrepancies, it is still clear that PPVs are low in low-prevalence populations.

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*Human T-lymphotropic virus type I (HTLV-I) Dupont HTLV-I ELISA: Product Insert Sheet

Our criteria for seropositivity have been recommended by the US Public Health Service Working Group[10] and require reactivity to both *gag* and *env* gene products; they currently necessitate at least Western blotting and in some instances WIB and RIPA as *env* reactivity appears to be better on RIPA than on WIB. The significance of indeterminate HTLV-I serologic results is not fully understood; these may represent nonspecific reactivity to cellular proteins, cross-reactivity with other unknown retroviruses, or early seroconversion to HTLV-I. The uncertainty regarding the significance of HTLV-I indeterminates resembles the situation with human immunodeficiency virus (HIV) where the full significance of sera with "indeterminate" HIV WIB remains to be determined[16]; further studies including serial serologic testing of persons with indeterminate HTLV-I serologies and examination of peripheral mononuclear cells with the new polymerase chain reaction (PCR) technique,[17] we hope, will provide evidence for or against viral infection. In analyzing our data, we excluded the indeterminates and thus avoided the issue of whether to count them as positives or negatives. Including the indeterminates with the positives would lower the calculated sensitivities of the screening tests; including them with the negatives would lower their specificities. Either way the resulting PPVs would be lower than the ones we calculated. For example, at a prevalence rate of 0.1 percent, including the indeterminates with the positives produces a PPV of 4.5 percent for Dupont's (compared with 7.3 percent) and 2.0 percent for Cellular Products' (compared with 2.3 percent); including them with the negatives changes the PPVs to 4.1 percent and 1.0 percent, respectively.

Based on our data, 11 to 21 percent of repeatably reactive specimens on EIA may be negative for HTLV-I antibodies (no bands on WIB). As blood is discarded from donors whose EIA results are repeatably reactive after screening, as much as 21 percent of discarded blood may be HTLV-I-negative. However, the objective of screening is an HTLV-I-free blood supply because HTLV-I is transmissible by transfusion of HTLV-I infected blood[18]; screening of the blood supply should achieve this. A limitation of current HTLV-I serologic tests, including WIBs and RIPAs, is that they cannot distinguish infection with HTLV-I from infection with HTLV-II, a closely related human retrovirus with no clear disease associations.[19] At present, HTLV-I and HTLV-II can be separated by amplification of viral specific sequences in infected cells using polymerase chain reaction (PCR).[20]

Acronyms Used

ATL, adult T-cell leukemia/lymphoma; CDC, Centers for Disease Control; EIA or ELISA, enzyme-linked immunoabsorbent assay; FDA, Food and Drug Administration; HAM, HTLV-I-associated myelopathy; HIV, human immunodeficiency virus; HTLV, human T lymphotronic virus (I and II); PAGE, polyacrylamide gel; PCR, polymerase chain reaction; PPV, positive predictive value; RIPA, radioimmunoprecipitation assay; TSP, tropical spastic paraperesis; WIB, Western immunoblot.

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Public Health Briefs

Comparison of Two Methodologies to Measure Agricultural Occupational Fatalities

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Abstract: Agricultural occupational fatalities in Pennsylvania for the years 1985–87 were followed up. Supplemental data concerning the occupation of the deceased and circumstances of the fatal accident were obtained from a family member. The number of fatalities designated as agriculturally and occupationally related by the National Traumatic Occupational Fatality (NTOF) system was compared with the number of fatalities identified by using alternative criteria for classification of agriculturally and occupationally related fatalities. There may be nearly a 30 percent error in the NTOF method resulting in a 20 percent undercount. *(Am J Public Health* 1989; 80:198–200.)

Introduction

The National Safety Council (NSC) estimates that agricultural industry fatalities for 1984* were 46 deaths per 100,000 workers[1] while the National Institute for Occupational Safety and Health (NIOSH) estimates them at 18.36 for the same year.[2]

The reliability and validity of the data used to generate these estimates have been questioned. Of specific concern is that these data are derived from information on death certificates regarding usual occupation, type of business or industry, and injury at work categories.[2–6] In agriculture, occupational work is difficult to distinguish from non-occupational work since the worksite is often the place of residence. This combination results in exposure to worksite hazards in both working and non-working situations and in both occupational and non-occupational activity.[7–9]

Additionally, the agricultural labor force is composed in part of individuals less than 16 years of age, individuals who work in other occupations on a full-time basis, and family members and relatives who may lend a helping hand only occasionally. These labor characteristics are not normally the case in other industries. Furthermore, agricultural land is often used for recreational purposes such as hunting, swimming, riding, etc, and for both non-occupational and occupational work. Lastly, differing methodologies are used to record and classify agricultural fatalities by the NSC and NIOSH.[1,2]

This paper compares the number of fatalities designated as both agricultural and occupational by the National Traumatic Occupational Fatality (NTOF) system, a system developed by NIOSH, with the number of fatalities identified by using alternative criteria for classification.

Methods

NIOSH developed the NTOF classification system to monitor occupational fatalities for all industrial sectors in the United States, including agriculture. The NTOF system collects and analyzes data on traumatic occupational fatalities using information from death certificates recorded by vital statistics reporting units in all 50 states and the District of Columbia.[2] These selected death certificates are then coded for the "usual industry" of the victim by divisional level industry categories listed in the Standard Industrial Classification Manual.[2,10] The assumption is made that "usual industry" at the divisional level is a proxy for the occupation engaged in at the time of the fatality. The number of death certificates coded into the agricultural industry are then considered agricultural occupational fatality cases.

We obtained copies of all death certificates for the years 1985–87 that indicated deaths that were accidental, and included "farm" or "agriculture" in any category (N = 141) from the Pennsylvania Department of Health, Division of Health Statistics and Research. For the same time period, a clipping service forwarded all newspaper accounts of farm and agricultural accidents and fatalities occurring in Pennsylvania.

The sample of fatality cases was reduced to 100 since only the death certificates that could be matched with a newspaper clipping containing the name and address of the deceased member's family were used. Of these 100 fatalities, 66 (66 percent) individuals from families of the deceased agreed to participate in the study. This group will hereafter be referred to as the supplemental data group.

A family member, usually a spouse, parent, or child of the deceased, was contacted by either telephone or mail to obtain supplemental data on the deceased's primary or usual occupation, secondary or part-time occupation, and the relationship between the activity at the time of the fatality and the deceased's primary or secondary occupation, non-occupational work, recreational pursuits, or other activities. The terms "primary or usual occupation" and "secondary or part-time occupation" were not further defined to the individuals questioned. Further information on the methodology

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*The agricultural industry, as defined by Standard Industrial Classification (SIC) codes 01 through 09 includes production agriculture (farms, ranches, commercial orchards, truck gardens, etc.), various agricultural services, commercial hunting and fishing, and forestry, excluding logging.

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used to collect the supplemental data are detailed elsewhere.**

The alternative criteria for designating an agricultural occupational fatality included cases whose victims were:

- engaged in a work activity related to *either* their primary or secondary occupation of agriculture at the time of the fatality,
- under the age of 16 and engaged in a farm work activity, or
- under the age of 16 and injured by exposure to an agricultural worksite hazard although not engaged in an agricultural work activity.

Results

Our alternative criteria and the supplemental data were used as one data source when comparisons, explained below, were made. Because of the single data source, the effect of the alternative criteria and the supplemental information cannot be separated in this analysis. Table 1 compares characteristics of three groups: cases with supplemental data, cases that did not respond to our survey, and cases for which there was no match with a newspaper clipping. There are no differences between the 66 cases followed up and the other two groups.

Twenty-six of the 66 cases (39 percent) were identified by the NTOF criteria as agricultural occupational fatalities; the remaining 40 cases (61 percent) were excluded. Using the alternative criteria and supplemental information, 39 of the 66 cases (59 percent) were classified as agricultural occupational fatalities; the remaining 27 cases (41 percent) were excluded.

TABLE 1—A Comparison of Selected Variables between the Supplemental Data Group and the Non-response and Death Cerficate Only Groups

	Groups		
Variables Compared	Supplemental Data Group (%)	Non-response Group (%)	Death Certificate Only Group (%)
Sex			
Male	63 (95)	33 (97)	35 (85)
Female	3 (5)	1 (3)	6 (15)
Age (years)			
Under 16	9 (14)	8 (24)	7 (17)
16 or over	57 (86)	26 (76)	34 (83)
Time of death			
AM	22 (33)	6 (18)	9 (22)
PM	44 (67)	27 (79)	31 (76)
Unknown	—	1 (3)	1 (2)
Injury at work			
Yes	31 (47)	19 (56)	13 (32)
No	28 (42)	14 (41)	25 (61)
Blank	7 (11)	1 (3)	3 (7)
Accidental injury			
Yes	63 (95)	31 (91)	40 (98)
No	—	1 (3)	—
Blank	3 (5)	2 (6)	1 (2)
Medical examiner			
Yes	56 (85)	30 (88)	35 (85)
No	4 (6)	2 (6)	3 (7)
Blank	6 (9)	2 (6)	3 (7)
Autopsy			
Yes	22 (33)	12 (35)	16 (39)
No	42 (64)	20 (59)	24 (59)
Blank	2 (3)	2 (6)	1 (2)

Of the 26 cases designated by NTOF as agricultural occupational fatalities, 3 (12 percent) were found to be non-agricultural occupational; the misclassification was due to relying on "usual industry" as a proxy for the occupation engaged in at the time of the fatality.

Sixteen of the 40 cases (40 percent) excluded by the NTOF criteria should have been classified as agricultural occupational fatalities. There were three reasons why these cases were missed by NTOF:

- In three cases, the "injury at work" category on the death certificate was either marked "no" or left blank.
- In seven cases, the fatality was from the secondary occupation of agriculture and secondary occupations are not listed on the death certificate.
- In six cases, the victim was under the age of 16 and was involved in farm work or was fatally injured from exposure to a worksite hazard.

Of the agricultural occupational cases NTOF did identify, three should *not* have been included. This means 19 of 66 cases were misclassified. However, because 16 cases were undercounted and three cases were overcounted, the net error was an undercount of 13 (20 percent) agricultural occupational fatalities.

The 27 cases that were excluded by our alternative criteria and supplemental data were examined to determine the specific reasons for their exclusion. These cases fell into four groups (Table 2) and are clearly non-agricultural and/or non-occupational.

Discussion

Our research suggests that there may be nearly a 30 percent error in the NTOF method for designating agricultural occupational fatalities which resulted in a 20 percent undercount of these fatalities. This error is due to problems of incomplete and inaccurate information on the death certificates, NTOF's limiting criteria for inclusion as an agricultural occupational fatality case, and the use of "usual industry" as a proxy for work engaged in at the time of the fatality.

Adjusting the NIOSH estimate (18.36 per 100,000 workers) upwards by 20 percent still leaves their estimate well short of the National Safety Council estimate (46 per 100,000 workers). There are at least two possible explanations for the remaining difference. One reason may be the different denominators each used. NIOSH used Census Bureau data for their denominator,[2] while the NSC used Bureau of Labor Statistics data.[1]

A second reason is that the NSC estimate may include some cases similar to the 27 cases that were excluded by our alternative criteria and supplemental data (we found 15 cases that involved non-occupational work). Death certificates do not differentiate

TABLE 2—Supplemental Data Group Identification of the 27 Non-agricultural Occupational Fatalities

	Number
Non-Occupational Work	15
Recreational	4
Underlying Medical Condition	5
Operator/Passenger in a Licensed Motor Vehicle*	3
Total	27

*A fatality where the victim was an operator or passenger in a licensed motor vehicle and was involved in a collision with agricultural machinery, livestock, or other agricultural hazard. A public roadway accident is the primary example.

**Seltzer B, Murphy D, Yesalis C: A methodology for the collection of supplemental information on agricultural fatalities. Paper submitted for publication, September 1988.

between occupational and non-occupational work, but unless supplemental data are collected, it is difficult to differentiate these fatalities. Differentiating fatal work accidents that are both agriculturally and occupationally related from those that are simply work related or that happen to occur on a farm or in a rural area is critical if agriculture, as an occupation, is to be fairly and accurately compared with other occupations.

The supplemental data collected in this study have additional benefits besides helping to more accurately identify agricultural occupational fatalities. This information can also yield information on non-occupational related injury and health problems. The Cooperative Extension, which has responsibility for farm and home safety education, can use this information to guide safety education and injury prevention activities. Yet another benefit is to assist state agencies to appropriately allocate resources for injury prevention.

The agricultural industry's unique workforce and workplace characteristics pose a greater challenge than most other industries for the tracking of its occupational work fatalities. The NTOF system is a good start for enumerating fatal agricultural occupational work cases at the national level and it is replicable for a single industry at the state level. This allows those interested in single industry state level data to conduct follow-up studies such as we have reported. These state studies can be used as a basis for state statistics, and for adjusting particular industry totals at the national level.

This study should be replicated in several states to better quantify errors of both the NTOF system for agriculture, and non-occupational fatal injuries in agriculture.

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Use of OSHA Inspections Data for Fatal Occupational Injury Surveillance in New Jersey

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Abstract: Occupational Safety and Health Administration (OSHA) computerized inspections data, death certificates, and medical examiner records identified 204 fatal occupational injuries in New Jersey, 1984–85. OSHA computerized data uniquely identified seven cases. They did not identify 35 fatalities under OSHA's jurisdiction, of which 24 were investigated by OSHA but not recorded, four were not considered work-related, and seven were not known to OSHA. Eighty-seven were outside OSHA's jurisdiction; 28 were among the self-employed who are not under the health and safety protection of any governmental agency. (*Am J Public Health* 1990; 80:200–202.)

Introduction

Surveillance data for fatal occupational injuries have been compiled in several states.[1–5] These data systems are based largely on death certificates, but are usually supplemented by one or more sources of data such as workers' compensation reports and medical examiner reports. Each source independently captures some fraction of cases. Another potential source of data for surveillance systems is data from occupational fatality investigations carried out by the Occupational Safety and Health Administration (OSHA). Investigation data are maintained in a computer database, OSHA's Integrated Management Information System (IMIS).

We examined the usefulness and completeness of these OSHA inspection data for New Jersey's fatal occupational injury surveillance system. The objectives were to determine: whether OSHA investigation data identified cases not captured on death certificates or medical examiners' records; and whether OSHA investigation data were available for all work-related deaths within OSHA jurisdiction that were identified from other data sources in the surveillance system.

Methods

Unintentional occupational fatalities were ascertained for calendar years 1984–85. A fatal occupational injury was defined as a death resulting from an unintentional workplace injury. New Jersey cases were identified from the New Jersey computerized death certificate file by a code that indicated unintentional injury at work and from medical examiner records by manually sorting through reports in the state medical examiner's office. Cases from each source were then merged into one computerized file by matching on the name of deceased.

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The New Jersey data were compared to OSHA fatality investigation data identified initially from the OSHA IMIS file. The OSHA IMIS file is intended to contain reports of all workplace investigations conducted by OSHA area office inspectors. Name of deceased, name of employer, and other investigative information are included in each record. Names of deceased in the IMIS file were matched with names from the merged death certificate/medical examiner file.

Next, a list of decedents was compiled from cases in the death certificate/medical examiner file for whom a record was not matched in the IMIS file, but whose workplaces appeared to be under OSHA jurisdiction. Jurisdiction was determined by reviewing information about each decedent's workplace available on the death certificate and medical examiner report. Employers in the public sector are not under OSHA jurisdiction. Almost all employers in the private sector are covered by OSHA, except where other federal agencies including the US Department of Transportation and the Federal Aviation Administration have jurisdiction.

The list of decedents' workplaces potentially under OSHA jurisdiction was sent to the OSHA regional office to determine whether these employers were, in fact, under OSHA jurisdiction. If so, the regional office was asked to determine whether an investigation was carried out that was not recorded in the IMIS computer file.

Results

OSHA IMIS data listed 77 fatalities for 1984–85. Sixty-three (82 percent) of these matched with cases in the medical examiner/death certificate file. Four of the remaining 14 cases were cardiovascular deaths and therefore did not meet the case definition. No death certificate could be found for three of the 14. This could have been due to name mispelling or to coding errors in the IMIS file. The remaining seven of the 14 met the case definition and were not captured in the New Jersey data sources (Table 1). Death certificates were obtained for these seven cases; coding errors explained why they had not been captured.

Two-hundred and four fatalities were counted by combining the IMIS data with the death certificate/medical examiner file. Table 2 shows OSHA jurisdiction of the 204 cases. Eighty-seven (43 percent) of the fatalities were outside of OSHA jurisdiction. Of the 105 under OSHA jurisdiction, OSHA IMIS identified 70 (67 percent) cases. Twenty-four investigations were carried out by OSHA area offices but were not recorded in the IMIS system. Of the remaining 11 fatalities under OSHA jurisdiction, OSHA claimed that in four cases the deaths were not work-related because the decedents had preexisting medical conditions (e.g., seizure disorder) related to the fatal episode, and therefore an OSHA investigation was not required. Seven cases of the 105 under OSHA jurisdiction were not investigated because OSHA was unaware of the fatality.

TABLE 1—Occupational Fatalities Identified in OSHA IMIS Data: 1984–85

	Number	%
Matches with death certificate/medical examiner file	63	81
False positives (cardiovascular deaths)	4	5
No death certificate found	3	4
True positives not matching death certificates/medical examiner file	7	9
Total cases in IMIS data file	77	100

TABLE 2—OSHA Jurisdiction for 204 Occupational Fatalities Identified in New Jersey 1984–85

Cases	Number (%)
Not under OSHA Jurisdiction	
Transportation	51
Government	8
Self-employed	28
Total not under OSHA jurisdiction	87 (43)
Under OSHA Jurisdiction	
Investigated according to IMIS data	70
Investigated according to OSHA regional office, not in IMIS	24
Not work related, according to OSHA Regional office	4
Not investigated; OSHA regional office unaware of fatality	7
Total under OSHA jurisdiction	105 (51)
Insufficient Data to Determine OSHA Jurisdiction	12 (6)
Total fatalities	204 (100)

Discussion

This study showed that the OSHA IMIS data file identified occupational deaths identified through no other source, but that it did not serve as a sole source of information because it did not capture all work-related deaths. Because IMIS data capture cases independent of other data sources, these data should be included in state-based fatal occupational injury surveillance systems. In addition, these data could be used as another source to estimate the true magnitude of fatal occupational injuries nationwide, following the procedures of Suruda.[6] However, there are several problems in utilizing these data, including the failure of IMIS data to accurately reflect the numbers of fatality investigations completed by OSHA area offices and underascertainment due to limited OSHA jurisdiction. The national OSHA Office of Data Analysis is addressing the first problem.

In a previous report, with incomplete case follow-up by the OSHA regional office, it appeared that as many as 34 percent of cases with employers apparently under OSHA jurisdiction were not investigated.[*] Other states using data on employers' fatality reports to OSHA had similar findings.[2,5] After exhaustive examination of case files by the OSHA regional office, it has been determined that a smaller proportion (7 percent) of cases under OSHA jurisdiction were not actually investigated by OSHA. Currently, death certificate and medical examiner data in New Jersey are being shared with the OSHA regional office to ensure that all occupational fatalities within OSHA's jurisdiction are investigated.

These data showed that nearly half of these New Jersey occupational deaths were outside of OSHA jurisdiction, of which 28 were among the self-employed, who are not under the health and safety protection of any governmental agency. Investigative data from OSHA have certain details about the circumstances of fatalities that are not in data from other sources. Such details, which are important for injury prevention programs, are not available on these cases. Legislative changes are needed to address the large number of occupational deaths not under OSHA's jurisdiction.

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[*] Stanbury MJ, Goldoft M, O'Leary K: Data Sources for a Fatal Occupational Injury Surveillance System in New Jersey. Paper presented at the Annual Conference of the Society for Occupational and Environmental Health, April 6–8, 1987.

ACKNOWLEDGMENTS

Dr. Goldoft compiled the death certificate and medical examiners' data while serving as an Epidemic Intelligence Service officer with the Centers for Disease Control. Opinions expressed in this paper are those of the authors and do not necessarily reflect the views of the Centers for Disease Control. We wish to thank Kathleen O'Leary, Allison Tepper, and Kenneth Rosenman for their helpful review of previous drafts of this paper, and also the staff in OSHA for their cooperation in following up on fatality inspection records.

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Workplace Health Hazards: Analysis of Hotline Calls Over a Six-Year Period

JULIA QUINT, PHD, MARGARET HANDLEY, BA, AND KATE CUMMINGS, MPH

Abstract: Between 1981–1986 a state-based occupational health telephone hotline received more than 8,000 inquiries on over 3,000 hazardous agents. Major caller groups were employees (37%), employers (20%), health care providers, primarily physicians (19%), government agencies (12%), and labor unions (6%). Employees were the fastest growing caller group. Callers inquired about general health hazards of chemicals (65%), the relation of symptoms to work (22%), and risks to pregnancy (13%). (*Am J Public Health* 1990; 80:202–204.)

Introduction

In 1979, the California legislature enacted a right-to-know law with several support resources including the Hazard Evaluation System and Information Service in the Department of Health Services. As described elsewhere,[1] staff (toxicologists, physicians, industrial hygienists) perform a variety of functions, including operation of a publicly accessible telephone inquiry response system. Over 12,000 inquiries involving a cross-section of California workplaces have been received from 1980 through early 1989.

We present and discuss descriptive statistics which characterize inquiries received from 1981 through 1986.

Methods

Inquiry Intake Procedure

Information in the following categories is obtained from callers and recorded on a standard form after assuring confidentiality. Callers' concerns are grouped into one of three areas: effect on pregnancy; relation of symptoms to work; general health hazard information. Only one concern is recorded as primary; the order of priority (highest to lowest) is pregnancy, symptoms, general information. Callers are categorized as follows:

- *Employees:* persons calling on their own behalf regarding exposures in their workplace, self-employed individuals, and friends and relatives who call on behalf of employees.
- *Employers:* supervisors and managers, company health and safety and medical personnel and industrial hygenists.
- *Health care providers:* physicians, nurses, genetics counselors, poison control center staff and paramedics.
- *Governmental representatives:* California Occupational Safety and Health Administration (OSHA) personnel, city, county, state and federal personnel.
- *Others:* attorneys, news media representatives, volunteer organizations, educational institutions, etc.

Callers' characteristics and primary concerns were tallied from 1981 through 1986. Primary concerns by caller group were tallied from 1986 logs. Data on employee industries, occupations and agents and repeat callers were tallied from the 4th quarter, 1986. Industry data were available for 466 of 562 total inquiries.

Results

During the first four years of full service, there was a modest increase in the annual volume but in 1985 and 1986, the number of inquiries increased by 49 percent and 43 percent, respectively (Figure 1). The average number of inquiries during 1985 and 1986 (2,067) was approximately double the average (1,038) during 1981–84. The sudden increase in inquiry volume appears to be due to increased awareness regarding the potential health effects of chemical exposures at work.

In 1986, of the 365 employer inquiries received, 40 percent were from managers and supervisors, 29 percent were from health and safety personnel (other than industrial hygienists), and 24 percent were from industrial hygienists.

Of the health care providers who called in 1986, 67 percent were physicians (70 percent from private practices). Nurses (12 percent) and genetics counselors (9 percent) were the other major groups who called in 1986.

Approximately one-third of the government agency inquiries in 1986 were from California OSHA. The majority of these callers were industrial hygienists requesting information for worksite inspections.

Analysis of a subset of 1986 callers (562 inquiries received in the fourth quarter of 1986) showed that 39 percent of the employers, 36 percent of the health care providers, 33 percent

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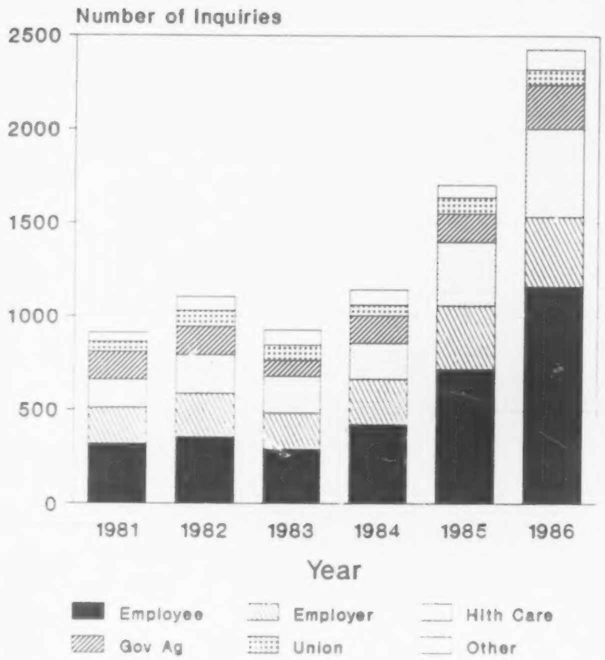

FIGURE 1—Number of Telephone Inquiry Response System Inquiries by Caller Group, 1981–86 (Caller groups defined under Methods)

of the union representatives, and 26 percent of the callers from government agencies were repeat callers. By comparison, only 2 percent of the employees had used the hotline previously.

Based on 466 inquiries from all callers in the fourth quarter of 1986, 48 percent concerned employment in service industries, 28 percent in manufacturing industries, and 8 percent in construction industries. The remaining inquiries concerned employment in transportation, communication, electric, gas, sanitary services (6 percent), agriculture (4 percent), public administration (4 percent), and retail (2 percent) industries. As a point of reference, the service industry was about 24 percent and the manufacturing industry about 18 percent of all non-agricultural California employees in 1986.[2] The types of industries included under services and manufacturing, along with typical occupations and agents of concern, are described in Tables 1 and 2, respectively.

Each year, multiple (10 or more) inquiries are received on 40 to 50 agents. The top 40 agents for which consultations were provided in 1986 are listed in Table 3.

Most caller concerns fall into three major categories. Average percentages of total inquiries for each category from 1981 through 1986 were 65 percent for general health hazard information, 22 percent symptom-related, and 13 percent pregnancy outcome.

The number of pregnancy-related inquiries received between 1981 and 1986 varied greatly. From 1981 through 1983 an average of 56 inquiries were received per year compared to an average of 375 inquiries per year from 1984 through 1986. A complete analysis of the pregnancy inquiries will be published in a separate report.[3]

Further analysis of 584 symptom-related inquiries received in 1986 showed that most were from employees (51 percent) and health care providers (26 percent), mainly physicians. Small percentages of symptom inquiries were from employers (8 percent) and government agencies (7 percent). California OSHA industrial hygienists inquiring about symptomatic employees encountered during worksite inspections represented about 60 percent of the symptom-related inquiries from government agencies.

Discussion

The preceding statistics clearly demonstrate that there is an increasing demand for acute and chronic health effects

TABLE 1—Service Industries,[a] Typical Occupations and Agents Inquired about during 4th Quarter, 1986

Industry	Occupation	Agents
Business (N = 86)[b,c]	Office workers	Video display terminals (VDTs), carbonless copy paper, photo copier chemicals, paint, carpet glues, indoor air contaminants, pesticides, asbestos
	Indoor pesticide applicators	organophosphate pesticides
	Photographers, photo lab technicians	amines, formaldehyde, hydroquinone
	Custodians	alkalis, alcohols, cleaners containing glycol ethers
Personal (N = 34)	Cosmetologists	hair dyes
	Manicurists	acrylic nail products (acrylates)
	Dry cleaners	perchloroethylene
	Laundry workers	alkalis (various products)
Health (N = 32)	Dental lab technicians	mercury, epoxy resins
	Histology, pathology lab technicians	formaldehyde, xylene, alcohols
	Hospital lab technicians	biological hazards, radiation
	Dental assistants	nitrous oxide, methyl methacrylate, mercury, formaldehyde
	Nurses	antineoplastic drugs, ethylene oxide, organic solvents
	Anesthesiologists	nitrous oxide, anesthetic waste gases
Educational (N = 18)	Teachers	PCBs, carpet glues, indoor air contaminants, photocopier chemicals, asbestos, fluorescent lights
	Librarians	asbestos
Miscellaneous (N = 18)	Artists	lead, dyes, pigments
Miscellaneous Repair (N = 16)	Welders	welding fumes, metals
	Furniture refinishers	methanol, methyl ethyl ketone
Auto Repair (N = 9)	Mechanics	asbestos
	Car painters and strippers	methylene chloride, toluene, xylene, phenols, isocyanates (specialty paints)

[a] Coded using a combination of industry, occupation, and narrative information.
[b] N = Total inquiries.
[c] Total Service Industry Inquiries = 222. Industries with less than nine inquiries are not shown.

TABLE 2—Manufacturing Industries,[a] Typical Occupations and Agents Inquired about during 4th Quarter, 1986

Industry	Occupations	Agents
Electrical Machinery Equipment and Supplies (N = 29)[b,c]	Assemblers, Parts cleaners, Wafer fabricators/processors	soldering fumes, organic solvents, epoxy resins, methyl ethyl ketone, glycol ethers, arsine, fluorides
Transportation Equipment (N = 14)	Boat builders	plastics, fiberglass, ketones
	Ship builders	epoxy resins
Chemicals and Allied Products (N = 12)	Chemical workers	acrylamide, metals, sulfur compounds
Printing (N = 11)	Printers	methylene chloride, perchloroethylene, glycol ethers
	Silk screener	organic solvents, acids
Stone, Glass, Concrete Products (N = 11)	Glass factory employees	organotin, silica, kerosene
Fabricated Metal Products (N = 10)	Machinists	glycol ethers, organic solvents
	Welders	welding fumes
	Solderers	solder fumes, fluorocarbons
Machinery (Non-Electrical) (N = 7)	Machinists	lead, thorium, cadmium
Food and Kindred Products (N = 7)	Food processors	chlorine, phosgene, sulfur dioxide
	Winery employees	carbon dioxide, nitrogen compounds
Rubber and Misc. Plastic Products (N = 6)	Plastics/Teflon industry employees	organic solvents, polymers

[a] Coded using a combination of industry, occupation and narrative information

[b] N = Total inquiries

[c] Total Manufacturing Industry Inquiries = 130. Industries with less than six inquiries are not shown.

TABLE 3—Agents Most Frequently Inquired about in 1986

Agent	Number of Inquiries[a]
glycol ethers	119
methylene chloride	108
trichloroethane (1,1,1)	89
lead and lead salts	82
acetone	76
asbestos	74
formaldehyde	73
fluorocarbons	71
acrylates (including acrylic nails)	67
toluene	64
xylene	61
methyl ethyl ketone	57
solvents, not otherwise specified	54
isopropanol	49
polychlorinated biphenyls	48
ammonia	43
paints	43
video display terminals (VDTs)	41
malathion	39
ethylene oxide	39
methanol	38
isocyanates	37
mercury (organic, metallic, inorganic)	35
silica and silicates	33
chromium and chromium salts	33
sulfuric acid	32
perchloroethylene	32
hydrochloric acid	31
arsenic and arsenic salts	29
epoxies	27
chlorpyrifos	26
sodium hydroxide	24
monophenol	21
trichloroethylene	20
fluorescent light fixtures	19
chlorine	18
diazinon	18
nitric acid	17
benzene	17
carbon monoxide	17

[a] Number of times agent was inquired about. May not represent distinct consultations since one consultation often involved multiple agents.

information for workplace exposures. Provision of this information is the core of the Hazard Communication Standard.

Eighty percent of all employers calling in 1986 wanted general health hazard information (such as acute and chronic health effects and industrial hygiene information), suggesting that this group needs further assistance to meet fully the challenge of its responsibility.

Of the various callers, employees are the largest and fastest growing subset. In 1986, 38 percent of all employees were primarily concerned with general health hazard information while 36 percent called about pregnancy concerns and 26 percent called because they had symptoms. Calls from health care providers closely paralleled this distribution. Both employee and health care provider symptom calls focused largely on acute effects and required, for the most part, information and industrial hygiene interventions to reduce exposures.

The telephone hotline is sometimes a limited forum for disseminating useful information. Specifically, some callers need to be educated about basic principles of toxicology before their complex questions (e.g., what is the cancer risk?) can be answered.

The purpose of hotlines is to provide ready access to information rather than education. Perhaps the service would have broader impact if it targeted those individuals who can use and disseminate the technical information provided. Callers who would be potential targets for a technical assistance hotline (health care providers, government agencies, union representatives, and employers) are by virtue of their jobs in contact with many employees. Thus, the information provided to them may benefit many employees. The high percentage of repeat callers among these potential target groups indicates that the hotline, to some extent, has already established an identity as a technical assistance service.

Since hotlines do not necessarily present the best forum for some types of callers, we also publish and disseminate chemical fact sheets and other educational material.

ACKNOWLEDGMENTS

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Incentives and Competition in a Worksite Smoking Cessation Intervention

LEONARD A. JASON, PHD, SHYAMALANITA JAYARAJ, CARYN C. BLITZ, MARK H. MICHAELS, AND LORI E. KLETT

Abstract: We evaluated a worksite smoking cessation program that employed multicomponents including support groups, incentives, and competition. The combination of incentives and other components increased participation rates to over 80%. Forty-two percent of smokers were abstinent at six months, compared to 13 percent of a control group (difference 29 percent, 95% CI 9, 49). (*Am J Public Health* 1990; 80:205–206.)

Introduction

Klesges, Cigrang, and Glasgow[1] and Klesges and Cigrang[2] reviewed seven controlled worksite programs having incentives and/or competition. The average post-test cessation rate was 44.8 percent, and the follow-up cessation rate was 45.9% (if one study by Stachnik and Stoffeimayr[3] with an exceptionally high rate is excluded, the average 13-month follow-up rate was 26.1 percent).

In January of 1988, we began pilot work with a worksite incentive system, that built on our previous work.[4,a] We report the results below.

Methods

The intervention was composed of two parts: a three week quitting program, and six months of follow-up meetings. The two group leaders of the two groups were employees of the Experimental (intervention) company, who were provided training and consultation. One of the two groups met on company time in the late afternoon, the second group met during lunch. For the first three weeks, there were two meetings each week. For the next two weeks, there were weekly meetings. For the next four weeks, there were meetings every other week. For the last four months, there were group meetings on a monthly basis. Each meeting lasted from 40 to 50 minutes.

Incentives

Participants also could earn $10 for attending each meeting regardless of their smoking status. Following the end of the initial three-week program, participants could also earn $1 each day that they were abstinent up to $180 (six months), and for every 30 continuous days of abstinence an extra $30. At the end of the program, each participant who had been abstinent for at least the last two months was eligible to participate in a lottery, with the company contributing $50 for each person who quit in the program. (The company changed this policy at the end of the program by actually putting in $50 for each person who signed up for the smoking cessation program.)

Participants had the opportunity to compete in a team effort for additional incentives. Teams consisted of three members who worked together to quit smoking. The team with the largest number of days abstinent won a cash prize of $300 at the end of the program.

Participants also had the option of selecting a coworker "buddy" to aid in those times of temptation to start smoking again. At the program's end, each buddy was given an opportunity to select an item from a list of gifts.

Participants were not required to join a group. They could still participate and be eligible for other incentives. On a pre-intervention survey, 81 percent of Es indicated they wanted the group program, 4 percent only wanted the self-help book, 2 percent wanted to quit on their own, and 11 percent indicated that they might quit smoking. As for incentives, 91 percent requested payment at meetings, 94 percent payment of $1 a day for quitting, 98 percent for $30 for a month of continuous abstinence, and 93 percent for the lottery. Forty-five percent of Es selected a team and 57 percent selected a buddy. Participants earned an average of $237. Fifty-three percent of Es indicated they would not have tried to quit if the program had not been offered.

Control group smokers were not provided any program or incentives; the high rates of control participation can partially be explained by their interest in obtaining prepoint expired air carbon monoxide (CO) tests. In the CO test, participants first exhale completely, then take a deep breath, hold it for 20 seconds, expire half the air, then breathe into an impermeable plastic bag to provide a sample of carbon monoxide in their breath to be measured.[5]

Results

Pretest Comparability

Two companies agreed to participate in our worksite smoking cessation study. There were approximately 300 employees at the Experimental company and 260 at the Control company. Forty-two of the 52 Control company smokers and 53 of the 63 smokers in the experimental company agreed to be involved in this study.

The two populations were rather similar (Table 1). Experimentals had smoked longer than controls and had higher status jobs.

Program Impact

The experimental manipulation of support, incentives, and competition had a substantial effect on abstinence rates at the immediate post-test, six-month end-of-group meetings, and at the six-month follow-up (Table 2). Experimentals also had higher continuous abstinence rates and overall worksite quit rates (i.e., the percentage of all smokers, both participants and non-participants).

Most of the experimentals selected incentives and the support group, but several preferred to accept other incentives unrelated to group attendance. For two who only selected the self-help book plus incentives, one quit at the post-test and six-month end-of-group meeting, and both had quit by the 12-month follow-up. One who decided to quit on his own and also accept incentives was successful at the three

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[a] Jason LA, Lesowitz T, Michaels M, Blitz C, Victors L, Dean L, Kimball P: A worksite smoking cessation intervention involving the media and incentives. 1989. Manuscript under review, DePaul University, Chicago.

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TABLE 1—Demographic Characteristics and Preintervention Measures on Experimental and Control Conditions

	E	C	Difference (95% CI)
Gender			
% Male	20	22	2% (-16,20)
% Female	80	78	2% (-16,20)
Race			
% Black	8	11	3% (-9,15)
% White	92	89	3% (-9,15)
Occupation			
% Labor/Craft	37	42	5% (-17,27)
% Service	12	33	21% (3,39)
% Prof/Managerial	51	25	25% (4,46)
Education[a]	4.4	3.9	.5 (-0,1)
Age (mean years)	36.8	33.3	3.5 (-.4,7.4)
Mean years smoking	19.3	14.7	4.6 (.9,8.3)
Preintervention smoking			
Number of cigarettes per day	24.6	20.1	4.5 (-.8,9.8)
Tar (mg/cig)[b]	12.1	11.3	.8 (-1.3,2.9)
Nicotine (mg/cig)[b]	.81	.77	.04 (-.08,.16)
Carbon Monoxide (mg/cig)[b]	11.9	10.9	1.0 (-.8,2.8)
% Tried to quit before	82	67	15% (-5,35)
Mean number of prior quit attempts	3.4	3.1	.3 (-1.3,1.9)
% Previously stopped smoking 6 months or more	43	39	4% (-18,26)
Mean number of times stopped smoking 6 months or more	1.8	1.8	.0 (-1.2,1.2)

[a] A 7-point scale was used (Elementary School = 1, Professional School = 7)

[b] The participants told interviewers their brand of cigarette and these brands were checked with the January 1985 Federal Trade Commission Report to obtain tar, nicotine, and carbon monoxide levels.

E = Experimental
C = Control

TABLE 2—Immediate Post-test, 6-Month, and 12-Month Measures on Abstinence, Continuous Abstinence, and Worksite Smoking Reductions

	E	C	Difference (95% CI)
Post-Test (95)*	(53)	(42)	
% Abstinence	49	9	40% (21,59)
% Worksite Smoking Reductions	41	8	33% (17,49)
6-Month (92)*	(53)	(39)	
% Abstinence	42	13	29% (9,49)
% Continuous abstinence	34	5	29% (13,45)
% Worksite Smoking Reductions	35	10	25% (10,40)
12-Month (84)*[a]	(47)	(37)	
% Abstinence	36	16	20% (1,39)
% Continuous abstinence	21	5	16% (2,30)
% Worksite Smoking Reductions	32	13	19% (3,35)

*Total numbers in groups

[a] Of the 6 Es who left the company, 3 left abstinent; whereas more of the 5 Cs who left the company were abstinent.

time points. Of the six experimentals who indicated maybe they would quit but would also accept incentives, only one quit at the immediate post-test, relapsed at the end of six-month group meetings, and had quit again at the six month follow-up.

Biochemical Verification

CO confirmations (ppm ≤ 9) were conducted at both experimental and control settings at the post point and end of the six-month meeting period. There was an 85 percent accuracy rate with self-reports. When discrepancies occurred, there are a number of explanations that could have explained the findings (e.g., exposure to second hand smoke, poor ventilation at home or in the car).

Discussion

The present study indicates that incentives and competition can be used effectively to increase participation in worksite smoking cessation programs. Overall worksite reductions were high because a large percentage of smokers participated in the program. It should be noted that worksite policies that allowed smoking at designated areas were introduced in both settings prior to the start of the intervention. Some of the effectiveness of the present program might be due to the smoking restriction policies at the experimental company.

One of the advantages of the present study was that experimentals could select those parts of the program that most appealed to them. Although most chose the groups and incentives, several of those who quit selected other materials and/or incentives. This suggests that different individuals might prefer different types of programs.

There are some limitations to the present study. There were several prepoint differences between the experimental and control companies, and these differences might have played a role in the outcomes. In addition, only two companies were used, and thus this research can only be considered a case study until it is replicated with larger samples.

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Retirement and Primary Cardiac Arrest in Males

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Abstract: We investigated the association between retirement and primary cardiac arrest (PCA) in 126 male cases and controls, 25–75 years of age, without prior heart disease or comorbidity. After adjustment for age alone, retirement was not associated with an increased risk of PCA, (OR = 1.1; 95% confidence intervals = 0.5, 2.4). This lack of association was not uniform across age strata, however. In 10 of 19 discordant pairs 60 or more years of age, the control subject had been retired; in all seven discordant pairs under 60, the case had been retired (lower 95% CI of the relative risk = 1.9). (*Am J Public Health* 1989; 79:207–208.)

Introduction

Clinicians have long suspected that retirement might increase the risk of coronary heart disease (CHD).[1] Retirement represents one of several profound exit/loss events, such as widowhood, death of friends or relatives, divorce, and relocation, that become more prevalent with increasing age.[2] The literature on retirement and physical health is fairly small and marked by contradictory findings and important methodologic limitations.[3–5] We analyzed data from a community-based study of potential risk factors for primary cardiac arrest[6] to determine the influence of retirement. Since our study was restricted to persons who were clinically healthy, we were able to minimize potential bias related to the fact that poor health is one determinant of retirement.[7]

Methods

The design of this study has been presented in detail elsewhere.[6] Briefly, from reports by trained paramedics in Seattle and suburban King County, Washington between December 1979 and January 1981, all 1,250 cases of out-of-hospital primary cardiac arrest were identified. Cases were characterized by a sudden pulseless condition without evidence of a non-cardiac condition as the cause of cardiac arrest.

Because we required detailed information regarding the prior experience of each subject, we restricted the study population to married residents of King County, 25–75 years of age (779 cases). In addition, based on the emergency medical service report or a brief telephone screening interview with the spouses of potential cases, persons with the following conditions were excluded: a history of prior clinically recognized heart disease (492 cases), such as angina pectoris, myocardial infarction, coronary artery bypass surgery and graft, congestive heart failure, arrhythmias, cardiomyopathy, congenital or valvular heart disease; or comorbidity that "limited their usual activity" (103 cases), such as cancer, chronic lung, renal or liver disease. Thus, there were 184 cases eligible for inclusion in this study; they represented persons who appeared and acted as if they were "healthy" until the time of their primary cardiac arrest.

After each case was identified, a control subject matched for age (within seven years), sex, marital status, and urban or suburban residence was selected from the community using random digit dialing.[8] Spouses of potential controls were queried by telephone in a fashion similar to those of cases; potential controls with a history of prior clinically recognized heart disease or comorbidity were excluded. Of eligible subjects identified, the spouses of 89 percent of the cases and 85 percent of the controls complied with a more extensive in-home interview. Thus, there were 163 cases and 163 controls in our study population; 18 percent of subjects were female, 94 percent were White, and the mean age was 58 years.

The home interview verified the absence of prior clinically recognized heart disease or comorbidity and covered factors such as history of hypertension treatment, current cigarette smoking, habitual vigorous leisure-time activity, history of diabetes mellitus, and alcohol consumption. We also asked whether the subject was currently employed and, if not, whether the subject was retired and formerly employed, not working currently but temporarily laid off and seeking employment, or never employed. Information regarding the specific age of retirement and reasons for retirement was not obtained. These investigations were completed in 1981.

We restricted this analysis to the 127 male case-control pairs in which both members were either employed or retired. Among the six male case-control pairs that were excluded, four of the cases and two of the controls were temporarily laid off and seeking employment. One additional pair was excluded because of missing data related to employment status.

An unadjusted matched-pairs analysis and a conditional multiple logistic regression analysis for matched case-control data were used in this study.[9] Factors evaluated simultaneously using the matched-pair multiple logistic model included history of hypertension treatment, current cigarette smoking, habitual vigorous activity, and alcohol consumption. We included age as a continuous variable to control for residual confounding related to age. To explore possible differences between early and conventional retirement (as represented by age at interview), we used a dichotomous variable for age (<60 years, >60 years).

Results

Cases were more likely to be retired compared to controls (Table 1); comparing retired men with currently employed, the estimated relative risk was 1.7 (95% confidence intervals = 0.8, 3.4). However, the estimate of the relative risk was 1.1 either after adjustment for age alone (95% CI = 0.5, 2.4) or after simultaneous adjustment for age, history of hypertension treatment, current cigarette smoking, habitual physical activity, and alcohol consumption (95% CI = 0.5, 2.8).

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TABLE 1—Distribution of Primary Cardiac Arrest Case-Control Pairs According to Retirement Status, King County, Washington, December 1979–January 1981

Retirement by Control	Retirement by Case	
	Yes	No
Yes	27	12
No	20	68

Among case-control pairs less than 60 years of age, cases were more likely to be retired compared to controls (Table 2); in all seven discordant pairs it was the case who had been retired (relative risk = infinity, lower 95% CI = 1.9). Of note, among subjects less than age 60, the prevalence of cigarette smoking, obesity, diabetes, family history of heart disease, and alcohol consumption was similar among retired and employed men; however, prior hypertension treatment was more common among retired (62 percent) than employed (25 percent) men. There was no relation between retirement and the risk of primary cardiac arrest among men greater than or equal to 60 years of age, as reflected by an estimated relative risk of 0.9 (95% CI = 0.3, 2.7).

Discussion

There are several limitations of our analysis: we did not obtain specific information on the reasons for retirement; we do not know the extent to which retirement was involuntary. However, our study was restricted to persons without clinically recognized heart disease or comorbidity that limited their usual activity; only 8 percent of subjects had more than three restricted bed days over the prior six months and less than 1 percent had more than six days. For these reasons, we consider it unlikely that physical disability was a major determinant of retirement in our study. Since we did not obtain information on the time since retirement began, we were not able to determine if this factor influenced the risk related to retirement. Finally, because of the small number of retired men in our analysis, we were not able to fully explore potential confounding related to other risk factors for the age-specific estimates of risk related to retirement.

TABLE 2—Distribution of Primary Cardiac Arrest Case-Control Pairs* According to Age and Retirement Status, King County, Washington, December 1979–January 1981

Age	Retirement by Control	Retirement by case	
		Yes	No
<60	Yes	0	0
	No	7	57
>60	Yes	27	10
	No	9	4

*10 case control pairs where both subjects were not in the same age stratum were excluded from this table.

We chose a cut-point of age 60 for early retirement because we considered retirement prior to that age unambiguously early. For a later cut-point, such as age 62, the magnitude of the increase in risk related to early retirement decreased (relative risk = 2.8, 95% confidence intervals = 0.9, 8.2) but the direction of the interaction between age and retirement on the risk of primary cardiac arrest remained the same. Nevertheless, for the reasons noted above, our finding of an increase in the risk of primary cardiac arrest among younger but not older men should be viewed as preliminary. Because of the magnitude of the association of retirement with primary cardiac arrest that we observed in relatively young men, the benefits and risks of early retirement deserve further study.

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Seroconversion to Human Immunodeficiency Virus in Prison Inmates

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Abstract: We evaluated the prevalence and incidence of human immunodeficiency virus (HIV) infection in 3,837 inmates of a state prison system. Ninety-two (2.4 percent) were HIV-seropositive. The highest proportion of HIV-seropositive inmates was found among Blacks (5.4 percent), females, and those 30 years of age or older. HIV-seropositivity of entering inmates was also 2.4 percent and was unchanged over the three years of the study. Seroconversion occurred in two inmates while in prison, a rate of one conversion per 604 person-years, but HIV infection could have occurred before entry. Seroconversion to HIV was rare in inmates in this correctional facility (*Am J Public Health* 1990; 80:209–210.)

Introduction

Acquired immunodeficiency syndrome (AIDS) was first reported in prison inmates in 1983.[1,2] As of October 1, 1988, 3,136 AIDS cases had been reported in a survey of state and federal correctional facilities.[3] Since inmates frequently report intravenous (IV) drug use both before and during incarceration,[4,5] human immunodeficiency virus (HIV) infection in this population is not surprising. Studies of persons entering prison in the United States have shown rates of HIV seropositivity ranging from 0 to 17 percent.[3,6]

Several state prison systems, including the Nevada Department of Prisons, routinely screen incoming inmates for antibody to HIV. In Nevada, this is mandated by law. In August 1987, the Nevada Department of Prisons also began testing all inmates for antibody to HIV at the time of their release. To estimate how frequently HIV seroconversion occurs after inmates are admitted to a correctional facility and to assess whether seroconversion is the result of HIV exposure while in prison, we surveyed the results of HIV antibody testing of all inmates released from the Nevada Department of Prisons between August 1987 and July 1988.

Methods

The Nevada Department of Prisons operates 14 correctional facilities: one maximum and three medium security facilities for men, one women's facility, seven minimum security facilities, and two transitional centers. All persons admitted to the system beginning August 1985 received HIV antibody testing. In addition, all inmates in the system were tested in a prevalence survey from September through December 1985. Post-test counseling is offered to all inmates tested.

HIV antibody was determined by enzyme immunoassay (EIA; Electronucleonics Inc., Columbia MD, and Abbott Laboratories, Chicago, IL). Repeatedly EIA positive specimens were tested by the Western blot method at the University of California, Davis. Presence of an envelope band and a *gag* or *pol* band were required for Western blot confirmation.

Seropositivity of groups was compared by chi-square analysis. Confidence intervals were calculated using the Taylor series method.

Results

In September through December 1985, 92 (2.4 percent) of the 3,837 inmates* tested were HIV-seropositive. The highest proportion of seropositive inmates was found among Blacks (5.4 percent), females, and those 30 or more years of age (Table 1).

HIV seropositivity in incoming inmates from August 1985 to July 1988 was 124/5,216 (2.4 percent). In the first year of the study, the rate was 43/1,623 (2.6 percent); in year two, it was 43/1,758 (2.5 percent); and in year three, it was 38/1,835 (2.1 percent). No variation in age or race distribution of the affected inmates was seen over time (data not shown).

A total of 1,105 inmates had initial samples obtained on entry or in the prevalence survey and again on leaving the prison system. Of these 1,105 inmates, 1,066 (96.5 percent) were seronegative on initial testing and at exit; 36 (3.3 percent) were seropositive on both tests, two seroconverted, and one inmate was a possible seroconverter (Table 2). The seroconversion rate was therefore 2/1,069 or 0.19 percent of susceptibles. The 1,069 inmates whose initial test was negative were followed a total of 1,207 person-years; the seroconversion rate could therefore also be expressed as one conversion per 604 inmate-years.

To determine whether transmission had occurred within the prison system, we also assessed the HIV serostatus of all inmates in the system during the study period who had

TABLE 1—Prevalence of Antibody to Human Immunodeficiency Virus (HIV) in Nevada State Prison System Inmates, December 1985

Demographic Features	HIV Seropositive/ No. Tested (%)	Odds Ratio (95% CI)**	
Race			
White	21/2231 (0.9)		
All Other*	71/1606 (4.4)	4.87	(2.98, 7.96)
Sex			
Male	74/3594 (2.1)		
Female	18/243 (7.4)	3.81	(2.23, 6.48)
Age (years)			
Less than 30	32/1792 (1.8)		
30+	60/2045 (2.9)	1.66	(1.08, 2.57)

*Race	Number	(% Seropositive)
Black	1279	(5.4)
Hispanic	243	(0.8)
Native American	51	(0)
Other	33	(0)

**Kleinbaum DG, *et al*: Ref. 11

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*Racial-ethnic distribution of prison population: White 58.1 percent; Black, 33.3 percent; Hispanic 6.3 percent; Other 2.3 percent.

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TABLE 2—Seroconversion to HIV among Nevada State Prison System Inmates

Inmate Case	Age (years)	Sex	Race	Time in Custody before Testing	Risk Activity for HIV	Entry Test, Result	Exit Test Result
1	32	M	Black	130 days	Intravenous Drugs	EIA −	EIA +, WB +
2	35	M	Black	20 days	Intravenous Drugs, homosexuality	EIA −	EIA +, WB +
3	22	M	White	7 days	None known	EIA −	EIA +, WB indeterminate (p24 only)

entered on or before December 31, 1981; 313 such inmates were identified. Fifteen (4.8 percent) of these were HIV seropositive; however, all 15 had had their incarceration interrupted by parole at some time since 1982, and thus might have acquired HIV infection outside the prison.

Discussion

IV drug use has been cited as the major risk behavior for HIV seropositivity among prison inmates.[2,3] Our observation of a higher rate of HIV seropositivity in non-White inmates may therefore be a reflection of higher rates of IV drug use in non-White inmates or of higher rates of HIV infection in non-White IV drug users,[7] or both. The increased rate of HIV infection in incarcerated women in our survey may reflect an increased percentage of incarcerated women who are either IV drug users or sex partners of drug users. In Nevada, incarceration for a drug-related offense is more common among non-White than White inmates (18 percent versus 11 percent) and among female than male inmates (24 percent versus 13 percent); we therefore suspect that increased drug use in these groups accounts for the increased HIV seropositivity (Nevada Department of Prisons, unpublished data). Increased HIV seropositivity in older inmates may be related to duration of high-risk behavior for acquisition of HIV.

The rate of HIV seropositivity in incoming inmates (2.4 percent) remained stable over the three years of our study. Rates in incoming inmates in Maryland, although higher (7.2 percent), have also remained stable during the past three years.[8] The 2.4 percent seropositivity rate in incoming inmates in Nevada is higher than that reported in 20 of 25 prison systems responding to a recent Department of Justice survey.[3] Thus if HIV seroconversion within prisons such as those reported in the Department of Justice survey were occurring at a substantial rate, evidence of such seroconversion would have been seen in our survey.

Two definite cases of seroconversion to HIV were detected, while a third inmate developed only antibody to the HIV *gag* protein, p24, on Western blot. Reactivity to p24 alone is considered an "indeterminate" result but can be indicative of early seroconversion.[9] Because this inmate denied activity which would have put him at risk for HIV infection and refused retesting after his release, he could not be definitively classified.

The seroconversions could have resulted from exposure to HIV either before or after entering a correctional facility; although inmate 1 (Table 2) was in continuous custody for the four months before his first HIV antibody test, seroconversions occurring more than four months after exposure have been reported.[10] The overall rate of seroconversion (one case per 604 person-years) was even lower than the rate of one per 244 inmate-years reported from a study of inmates in the Maryland State Prison system.[6] However, in that study only a self-selected subset of exiting inmates were tested.

Both IV drug use and homosexual intercourse occur in prisons; in a study of inmates in Tennessee, 26 percent reported intravenous drug use and 17 percent reported homosexual activity while in prison.[5] Presumably, needle sharing and homosexual intercourse are also occurring among inmates in the Nevada Department of Prisons facilities. However, such occurrences are apparently not frequent enough between HIV-infected inmates and susceptibles to result in a high rate of HIV transmission. Because exposures could have occurred before entry into the Nevada State Prison System, the seroconversion rates reported here are maximum estimates of the transmission of HIV in prison inmates. We conclude that transmission of HIV among inmates incarcerated by the Nevada Department of Prisons is rare.

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Suicide Clusters: An Examination of Age-Specific Effects

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Abstract: The age specificity of time-space clusters of suicide was examined using National Center for Health Statistics data for 1978–84. Significant clustering of suicide occurred primarily among teenagers and young adults, with minimal effect beyond 24 years of age. Clustering was two to four times more common among adolescents and young adults than among other age groups. (*Am J Public Health* 1990; 80:211–212.)

Introduction

Approximately 29,000 people killed themselves in 1985 in the United States.* Suicide is rare before the age of 12, increases dramatically during adolescence, and reaches a peak between ages 20 to 24 at which time the suicide rate plateaus until another increase is witnessed among elderly men. Suicide among teenagers and young adults has increased markedly since 1960, moving suicide from the fifth leading cause of death to the second leading cause of death among teenagers in 1985.*

The recent occurrence of several highly publicized clusters of youth suicides has focused attention on the problem of suicide "epidemics"—unusually high numbers of suicides occurring in a small area and brief time period. Such outbreaks have been reported from ancient times through the twentieth century in diverse populations of varied ages,[1,2] but these reports are difficult to interpret without systematic surveillance and reporting of suicide clusters. Moreover, only recently[3] have statistical techniques for the detection of clusters of suicide in time and space been used, although several studies have examined temporal clustering after suicide stories.[1] Few studies have examined temporal fluctuations, in general[4–7] or spatial patterns.[7] Some clustering of suicides occurs by chance alone, even if suicides are occurring at random, so measurement of the rate of "outbreaks" must account for the clustering expected by chance variation.

Our earlier work indicated that clustering of suicides occurred among 15–19 year olds to a greater extent than would be expected by chance alone.[3] The present report examines the whole life span to determine whether suicide clusters are unique to adolescents or represent a more general phenomenon.

Methods

Data Sources

Vital statistics mortality data tapes 1978–84 were obtained from the National Center for Health Statistics (NCHS). Separate analyses were performed for the following age groups: 15–19, 20–24, 25–29, 30–34, 35–44, 45–54, 55–64, and 75+ years.

A suicide was defined as any individual whose cause of death had been listed as codes E950–959 of the Ninth Revision of the International Classification of Diseases (ICD-9). The late effects of a self-inflicted injury (E959) was included in the analyses because the timing of the death, in addition to the initial suicidal act, may influence a subsequent suicide. However, this classification of death only accounted for up to 0.2 percent of the total number of suicides within any age group and, therefore, it is unlikely that it had an impact on the analyses.

Statistical Method

The Knox procedure[8] was used to assess the degree of time-space clustering. Details on alternative procedures and the reasons for selection of the Knox procedure are given elsewhere.[3] The Knox procedure considers all possible pairs of cases and the time and space distances between them. Critical values for time and space to define "closeness" must be specified in advance. The test statistic involves a comparison of the observed number of pairs close in time and space with the expected number, which is the product of the proportion of pairs that are close in time with the number of pairs close in space. The test statistic approximately follows a normal distribution, $Z = (O-E)/\text{variance}$. The variance of the statistic was estimated following the method of David and Barton.[9]

Definition of Cluster

There are no explicit definitions of a "suicide cluster" in the literature. Initial operational definitions and limits for time and space parameters were examined in an exploratory analysis of NCHS data for 15–19 year olds for 1978 and state mortality data for 1978 through 1985, as described elsewhere.[3] Four space units (county and town of residence and occurrence) and five critical time units (3, 7, 14, 30 and 60 days) were examined to define closeness between pairs of cases. Exploratory analyses indicated that county of residence yielded a space unit more sensitive to the definition of a cluster. There was little difference between the five time units in the exploratory examination. The data are reported for 7, 14 and 30 days.

Results

As shown in Table 1, 1,420 close pairs of suicides (same county, 0–7 days) occurred among 15–19 year-olds, whereas 1,339.2 were expected. The 6 percent elevation above chance had an associated p-value of .011. The effect of clustering is primarily observed in the younger age groups (15–19 and 20–24 year olds) at the three critical time units. Very little clustering occurred at older ages, with the exception of the 55–64 year old age group at the seven-day critical time unit.

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*Mortality Statistics Branch, Division of Vital Statistics, National Center for Health Statistics. Vital Statistics of the United States (Unpublished data for 1985).

TABLE 1—Results of Knox Procedure on 1978–1984 NCHS Mortality Detail File

		Window = 7 Days				Window = 14 Days				Window = 30 Days			
		Close-Close Pairs				Close-Close Pairs				Close-Close Pairs			
Age (years)	N	Expected	Observed	O/E	P	Expected	Observed	O/E	P	Expected	Observed	O/E	P
15–19	12135	1339.2	1420	1.060	.011	2756.7	2842	1.031	.047	6000.7	6177	1.029	.009
20–24	25511	6335.5	6523	1.030	.008	13126.1	13481	1.027	.001	28642.6	29602	1.033	.000
25–29	23275	7126.1	7155	1.004	.365	14775.8	14888	1.008	.174	32247.7	32681	1.013	.008
30–34	19534	4805.0	4834	1.006	.338	9991.0	10035	1.004	.328	21794.4	21978	1.008	.106
35–44	29092	9707.5	9822	1.012	.119	20086.4	20271	1.009	.094	43780.9	44066	1.007	.087
45–54	26002	6976.1	7103	1.018	.060	14530.8	14629	1.007	.205	31622.9	31633	1.000	.478
55–64	25624	6399.6	6623	1.035	.002	13277.2	13401	1.009	.136	28928.5	29260	1.011	.024
65–74	19621	3334.8	3395	1.018	.139	6877.7	6976	1.014	.111	15002.1	15247	1.016	.020
75+	14718	2373.0	2367	0.997	.550	4876.8	4888	1.002	.435	10567.0	10554	.998	.551

The effect size was two to four times greater among the adolescents than in other age groups.

Discussion

Suicide clusters appear to occur primarily among teenagers and young adults, although clusters do occur occasionally at other ages. This is consistent with the findings of Phillips and Carstensen[10]: the impact of suicide stories on subsequent suicides was greatest for teenagers. In addition, although there is no formal surveillance system for suicide clusters in the United States, most of the suicide clusters reported to the Centers for Disease Control have primarily involved adolescents, young adults, or both.[1]

The magnitude of the clustering effect is relatively small, even among teenagers and young adults. Estimates of the percentage of teenage suicides that occur in clusters averages between 1 percent and 2 percent with considerable variation by state and year, yielding a range of estimates from less than 1 percent to 13 percent.[3] However, the relative risk for suicide given exposure to the suicide of one or more other persons may be quite great. Studies at the individual level will be needed to define this relative risk. Moreover, although the number of youth suicides that occur in clusters may be relatively small, they represent a class of suicides that may be particularly preventable. Suicide prevention efforts that focus on high-risk populations is a recommended strategy.[11]

The Centers for Disease Control has recently developed recommendations for a community plan for the prevention and containment of suicide clusters.[12] Targeting multiple risk groups in prevention strategies will be necessary to reduce the suicide rate substantially.

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Misclassification of Childhood Homicide on Death Certificates

GARRY D. LAPIDUS, PA-C, MPH, DAVID I. GREGORIO, PHD, MS, AND HOLGER HANSEN, MD, DRPH

Abstract: Suspect classification of homicide deaths of Connecticut residents under 20 years of age was noted for 29 percent of cases examined. Misclassification was attributed to incomplete or erroneous information recorded on the death certificates, rather than errors in the designation of ICD-9 homicide codes. The results have important implications in the interpretation of vital statistics when homicide is listed as the cause of death and underscore the value of record linkage systems. (*Am J Public Health* 1990; 80:213–214.)

Introduction

The causes of death reported on death certificates represent important national, state, and local statistics that are used in epidemiologic research as well as the planning/evaluation of health service delivery.[1,2]

Unfortunately, accuracy of death certificates regarding the cause of death presents problems.[3] In particular, under-recording of childhood homicide is known to be a serious problem.[4] It is thought to occur for a variety of reasons including ascertainment difficulties,[5] reporter unwillingness to consider homicide as a diagnosis,[6] and misclassification of homicides as "accidents" or as "cause of death unknown."[7]

This report addresses a complementary issue: the over-recording of childhood homicides on vital statistics.

Methods

Death certificates for all reported homicides between 1980–85 among Connecticut residents 0–19 years of age (N = 166) were reviewed by the authors. Supplementary information available from the State's Chief Medical Examiner Office on circumstances leading to the death of subjects (i.e., police investigation reports, eyewitness accounts and hospital records, laboratory and autopsy reports) was examined for the 161 deaths which occurred within the state.

The manner of death recorded on the death certificate (unintentional injury, homicide, suicide, or undetermined) was evaluated for consistency with supplementary information obtained from files of the State's chief medical examiners office. Homicide is defined in the ICD-9 CM manual as "injuries inflicted by another persons with intent to injure or kill by any means."[8] This definition is consistent with the legal interpretation of criminal homicide, i.e., "lethal violence (unlawfully) inflicted by a second person (or persons) with the intent to do bodily harm."[9] Non-criminal homicide deaths (i.e., negligence, justifiable homicide),[9] by contrast, would be excluded by a strict application of the ICD definition.

Certificates attributing death to the intentional use of a firearm, stabbing with a sharp instrument, other physical assault, arson and the like, when corroborated with statements by the assailant(s) or other eyewitnesses, and/or information from medical records and police investigations were considered accurate. Typical examples would be: an adolescent's death due to firearm injury where review of the medical examiner's report and police investigation report an argument following alcohol and/or illegal drug use in the company of friends or an acquaintance; a victim under three years of age for which a sworn confession by a parent attributes death to assault by blunt force was considered an appropriate designation of homicide.

Death certificates for which supplementary information was at odds were judged suspect for the purposes of this study. Our decisions were based on whether the police investigation documented criminal intent to injure or kill.

A preliminary review of circumstances leading to the death of cases suggested six causes of death for which an application of the ICD-9-CM homicide definition might be suspect: negligent use of a motor-vehicle, (e.g., motor vehicle/pedestrian collisions, without evidence of intent to injure, but rather hit and run incidents where the legal transgression was leaving the scene of a crash without stopping to give aid), unintended discharge of a firearm, (e.g., play with firearms by children, resulting in the unintended discharge of the weapon), suicide, legal intervention, arson, and drowning (e.g., where intent to harm was not documented).

Results

For 47 of 161 deaths studied (29 percent), we judged the manner of death reported on death certificates not to conform to strict interpretation of the ICD definition of homicide when supplementary information was taken into account. These included: 16 deaths due to arson, 13 motor vehicle pedestrian or pedal cycle collisions, 11 unintentional discharge of a firearm, four legal interventions, two suicides, and one drowning.

Table 1 characterizes those 47 cases. Error (20 percent) was less common regarding cause of death among 15–19 year olds, and most common (65 percent) among 10–14 year olds.

In addition, slight differences were noted by the race and sex of victim. Cause of death was misclassified in 17 of 74 cases among Whites (23 percent) compared to 30 of 87 among others (34 percent). Twenty seven percent of male cases compared to 33 percent of female cases were misclassified.

Discussion

Consistent use of the ICD-9 definition of death was found to be problematic in our investigation. Excluding such deaths from rate calculations would yield a corrected homicide rate of 2.13 per 100,000 persons, a reduction of some 30 percent of the current reported annual homicide rate for this age group.

We recognize the subjective nature of our effort and that additional information on the cases reviewed here (e.g., the legal disposition of criminal charges in such matters) may shed further light on the circumstances of intent to injure.

From the Department of Community Medicine and Health Care, University of Connecticut School of Medicine (all authors) and the Connecticut Coalition to Prevent Childhood Injuries (Lapidus, Hansen). Address reprint requests to Garry D. Lapidus, PA-C, MPH, Department of Community Medicine and Health Care, University of Connecticut Health Center, Farmington, CT 06032. This paper, submitted to the Journal December 23, 1988, was revised and accepted for publication June 23, 1989.

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TABLE 1—Suspect Homicide Deaths by Sex and Age of Victims; Connecticut Residents 0–19 Years, 1980–85

	Male Suspect cases							Female Suspect cases						
Age (years)	All Cases	Hit and run unintended injury	Playing with gun	Legal intervention	Suicide	Drowning	Arson	All Cases	Hit and run unintended injury	Playing with gun	Legal intervention	Suicide	Drowning	Arson
0–4	11		1				1	13	1					2
5–9	7					1	2	6		1				1
10–14	16	3	2	1			2	10	2	2				5
15–19	66	5	3	3	2		1	32	2	2				2
Total	100	8	6	4	2	1	6	61	5	5				10

47 of 161/29%

Furthermore, we recognize that underrecording of childhood homicides (i.e., false-negative cases) likely exceeds the discrepancies noted here. Our analysis focused exclusively on deaths for which the designation of homicide as a cause of death was problematic (i.e., false-positive cases). Attention to the complementary issue (i.e., the designation of criminal homicide deaths as unintentional injury, death due to other causes or cause of death unknown) was beyond our current resources but remains an area requiring further attention.

Our study pointed out several instances of failure to identify the cause of death correctly in circumstances involving the unintended use of firearms (e.g., playing with guns), hit and run motor vehicle collisions, legal intervention by police officers, arson, suicide and drowning. Patterns of misclassification related to the age of victim and mode of injury were detected. However, we were unable to specify systematic biases in the way data were recorded. Misclassification apparently resulted from incomplete or erroneous information on death certificates, rather than errors of ICD-9 homicide coding. Death certificates used for medical examiner cases require listing the manner of death as "accident, homicide, suicide, or undetermined." Presumably, initial homicide determination on the part of the medical examiner when forwarded to the state health department was then "properly" assigned an ICD-9 homicide code.

The importance of maintaining procedures for updating medical examiner files with related information available from police and hospital sources as well as revising death certificates on the basis of supplementary information that may be received is pertinent here. Permanent misclassification of deaths as homicides become more likely as procedures for revising certificates are inconsistent. In several instances we noted incomplete police reports. In others, supplementary information was dated after the certification of death date.

Problems stem, as well, from deficiencies of the ICD-9 classification system which do not allow accurate coding of "hit and run" motor vehicle pedestrian collisions. Roughly one-fourth (27 percent) of misclassified cases (13 of 47) were situations of this type. The coding of such cases as homicide rather than negligent use of a motor vehicle implies criminal intent by drivers. Such certificates provide little information to injury control researchers regarding the true circumstances of the event. Of the 161 cases reviewed only one true vehicular homicide was noted.

ACKNOWLEDGMENTS

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The Incidence of *Haemophilus Influenzae* Meningitis in a Midwestern Metropolitan County

RANDY B. RICE, BS, ANNE D. WALLING, MD, AND KEN J. KALLAIL, PHD

Abstract: The incidence of *H. influenzae* meningitis in children under five years of age was monitored during the years 1983–87 by a retrospective study of records from all hospitals serving a metropolitan county in Kansas. The mean annual incidence rate for the five years studied was 26.0 per 100,000. This compares with a mean of 56.0 per 100,000 measured in the same county during 1979–82. (*Am J Public Health* 1990–80:215–216.)

Introduction

Haemophilus influenzae type b (Hib) is the most frequent cause of bacterial meningitis in the United States,[1] accounting for 48.3 percent of all cases of bacterial meningitis[2] and causing an estimated 12,000 episodes of illness annually. Approximately 85 percent of *H. influenzae* meningitis cases occur in children below the age of five years.[1,2] Annual incidence rates show wide variation from 19 to 69 cases per 100,000 children under five years of age.[3] Of the six antigenic types of *H. influenzae*, type b accounts for 95 percent of the cases of meningitis.[4]

The purpose of this study was to describe the epidemiology of *H. influenzae* meningitis in children under age five in a midwestern community for the period 1983–87 and to compare the results with a similar study in the same community covering 1979–82.[5]

Methods

The study involved children residing in Sedgwick County, Kansas. The estimated total county population grew from 378,000 in 1983 to 394,000 in 1987 (Table 1). The number of children in the county under age five was estimated each year by the Metropolitan Long-Range Planning Department and updated by birth, mortality, and migration rates. In the under five population, the numbers of males and females were calculated to be equal and the racial composition was determined as 82 percent White, 11 percent Black, and 7 percent "other."

Hospital records were identified for all cases of *H. influenzae* meningitis in residents of Sedgwick County during the years 1983–87. The study included all hospitals serving the county population. Noncounty residents admitted to these hospitals were excluded from the study on the basis of zip code information. Cases missed by outflow to other medical centers are unlikely as the closest large population centers are more than 150 miles from Wichita. Examination of city/county health department records did not reveal any cases which could have died before hospital admission. Cases were identified by primary or secondary ICD codes and confirmed by culture and gram stain of cerebrospinal fluid.

Besides age-specific incidence rates, the variables measured included race, sex, and seasonal distribution. Data on *H. influenzae* meningitis in children under five years of age were extracted from the earlier investigation[5] which had studied meningeal infection of all etiologies in all age groups.

Results

Forty-one cases, all under age five, were identified. Forty were serologically type b and one was type d. These children accounted for all cases of *H. influenzae* meningitis in the community during the study period. The largest number of cases occurred between 12 and 17 months of age, followed by children younger than 6 months (Figure 1). Estimates of the population at risk were used to calculate annual incidence rates. As shown in Table 1, these ranged from 21.7 to 37.9 per 100,000 with a mean for the five years of 26.0 per 100,000 person-years.

H. influenzae meningitis occurred in 22 males and 19 females which corresponds to incidence rates of 27.9 per 100,000 for boys and 24.1 for girls during the five years studied. Thirty-six cases occurred in White children and five in Black (race-specific incidence rates were 27.8 for White children and 28.8 for Black). There were no cases in the "other" racial group. The differences by sex and by age were not statistically significant.

During the five years of the study, *H. influenzae* meningitis showed a bi-modal distribution with the highest number of cases occurring in late autumn and in the spring. The disease was least common in the summer months May through September.

Two White males with severe symptoms died within 48 hours of hospital admission during the study period. The case-fatality rate was 4.9 percent. The mean hospital stay for nonfatal cases was 12.2 days. At the time of hospital discharge, sequelae were recorded in 24 percent of hospital records. These sequelae included hearing loss, mental retardation, coma, acquired hydrocephalus, and respiratory impairments.

Comparison with Previous Study

Data from the earlier study[5] were recalculated to provide the incidence of *H. influenzae* meningitis in children under

TABLE 1—Incidence of *H. influenzae* Meningitis in Children below the Age of Five Years in Sedgwick County, Kansas

Year	Estimated Total Population	Population <5 Years of Age	Cases *H. influenzae* Meningitis	Hib Incidence in Children <5 years*
1983	378,000	30,890	7	22.7
1984	381,500	31,168	8	25.7
1985	387,800	31,667	12	37.9
1986	391,000	31,920	7	21.9
1987	396,000	32,239	7	21.7
Total	1,932,300	157,884	41	Mean = 26.0

*per 100,000 person years

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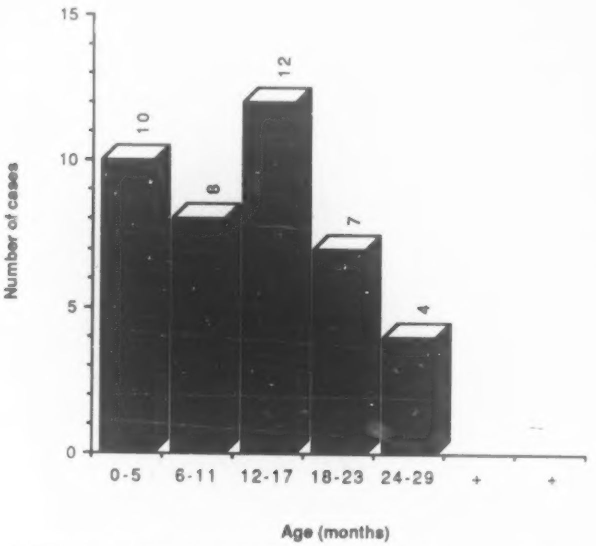

FIGURE 1—The Number of Cases by Age of *H. influenzae* meningitis in Sedgwick County, Kansas during 1983 to 1987

five years of age. The 65 cases observed over the four years (1979–82) corresponded to a mean annual incidence rate of 56 cases per 100,000 children under five, compared to a mean of 26.0 for the five years 1983–87, measured in the present study (95% confidence intervals of difference = 14,36). Methodology was similar for both studies.

Discussion

The assumption that all cases would be admitted to hospital in the county appears valid from our knowledge of clinical practices in the community. Outflow of cases to another city was unlikely as was the possibility of cases being treated without hospital admission. Cases could only have been missed through misdiagnosis and/or miscoding.

The most striking finding is the fall in incidence compared to the previous study. This could be due to several factors such as increased resistance of the population, lowered exposure to *H. influenzae*, decreased virulence of the organism, or the wide variations in annual incidence that occur naturally with this disease.[6] In view of the controversy over vaccination, it is tempting to attribute the fall to vaccination which was introduced to this community in 1986 and is generally offered at 24 months of age. It was not possible to identify data on the proportion of children immunized against *H. influenzae* in the county and information on vaccination was not recorded in the hospital charts. The age of the cases makes widespread immunization unlikely. Immunization could still be responsible for the fall in incidence by reducing cases in older children. This has the dual action of both directly reducing the number of cases and preventing infection being passed to vulnerable younger siblings.

Although the incidence of *H. influenzae* has fallen, the disease remains a potential cause of significant morbidity and mortality. The mortality rate of 4.9 percent confirms that reported by the Federal Centers for Disease Control.[1] One-fourth of all cases had evidence of serious sequelae at discharge and more pathology could be anticipated with long-term follow-up. In addition, the long length of stay represented considerable use of resources. While welcoming the fall in incidence, *H. influenzae* meningitis is a disease which commands respect and continued preventive efforts.

ACKNOWLEDGMENTS

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Trends in Hospitalized Discharge Rates for Head Injury in Maryland, 1979–86

ELLEN J. MACKENZIE, PHD, SHARON L. EDELSTEIN, SCM, AND JAMES P. FLYNN, MD, MPH

Abstract: Hospital discharge data from all acute care hospitals in Maryland were used to examine trends in hospitalized head injury incidence and outcome by severity. From 1979 to 1986, discharge rates increased by 3.4/100,000 per year; the largest percent increase was for more severe injuries. Discharge rates increased the most for adults ages 15–24 and ages 75+ but declined for children ages 0–4. Coinciding with the increase in head injury discharges was a decrease in the hospital case-fatality rate across all severity groups. (*Am J Public Health* 1990:80:217–219.)

Introduction

The public health importance of head injury as a major source of mortality and morbidity is well recognized. Recent studies have reported incidence rates that range from 152 to 200 per 100,000 population depending on case definitions and case ascertainment methodologies.[1–9] Despite differences in the magnitude of these estimates, findings regarding the patterns of incidence by age, race, sex, and severity are highly consistent. There is only one published report, however, in which these patterns have been examined over time (Olmstead County Minnesota, 1935–1974).[1] This paper describes the incidence of hospitalized head injury in Maryland over the eight-year period 1979–86.

Methods

Data Sources

Incidence rates of hospitalized head injury, 1979–86, were computed for the State of Maryland using uniform hospital discharge abstract data maintained by the State's Health Services Cost Review Commission for all 56 acute care, non-federal, hospitals as mandated by law. Excluded from the database are injuries treated on an outpatient basis as well as deaths that occur at the scene, in transport, or in the Emergency Department. Data abstracted include patient demographics, up to five ICD-9CM coded discharge diagnoses, up to three procedures (ICD-9CM coded), total and ICU length of stay, total and component charges, discharge status, and expected principal source of payment. There is a high degree of confidence in the quality of the discharge abstract data, including the diagnostic information, as they are used in paying all hospitals in the State.

To compute discharge rates the total number of hospital discharges (both residents and non-residents) was divided by the State resident population for that year as estimated by the Maryland Department of Health and Mental Hygiene. This approach assumes that the number of state residents treated in out-of-state hospitals roughly equals the number of non-residents treated in-state. All age-adjusted rates are based on direct adjustment using the 1980 census as the reference population.

Definitions

A head injury case was defined as any discharge with a principal or secondary diagnosis of a skull fracture and/or intracranial injury (ICD-9CM codes 800, 801, 803, 804, 850–854). A secondary head injury discharge was included only if the principal diagnosis was a traumatic injury to another body system; complications (ICD 9CM 958) and late effects (ICD 9CM 905–909) of injuries were excluded.

Injury severity was determined through the use of the software ICDMAP which maps ICD-9CM coded diagnoses into Abbreviated Injury Scale (AIS) scores. Studies have shown that AIS scores derived from ICDMAP are generally lower than those obtained by reviewing the entire medical record.[10] Discharges with more than one head injury diagnosis were classified according to the most severe (highest ICD/AIS score) head injury. Discharges with a head injury of ICD/AIS 1 or 2 are classified as minor, ICD/AIS 3 as moderate, 4 as serious, and 5 as severe.

TABLE 1—Age-Adjusted Head Injury Discharge Rates (per 100,000) by Year and Severity, Maryland, 1979–86

	Severity of Head Injury				
Year	Mild (ICD/AIS 1–2)	Moderate (ICD/AIS 3)	Serious (ICD/AIS 4)	Severe (ICD/AIS 5)	Total
1979	85.7	18.1	8.2	1.7	113.7
1980	82.1	18.3	8.4	2.9	111.7
1981	85.4	21.0	8.2	3.3	117.9
1982	87.6	20.3	8.3	3.6	119.8
1983	88.1	19.6	9.7	4.3	121.7
1984	92.5	20.3	11.1	5.0	128.9
1985	95.2	20.8	12.0	5.7	133.7
1986	94.3	23.1	11.8	5.0	134.2

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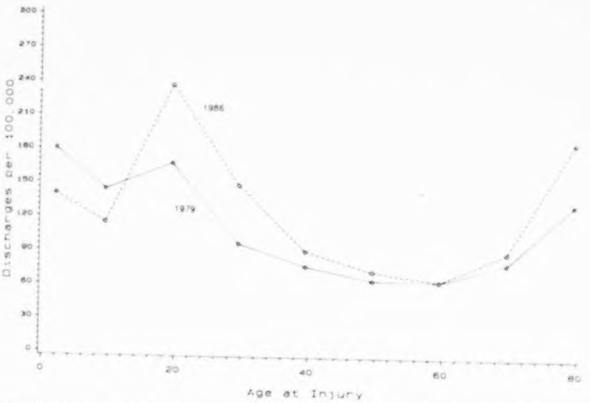

FIGURE 1—Head Injury Discharge Rates (per 100,000) by Age at Injury

Results

Since 1979 there has been an overall increase of 18 percent in the age-adjusted head injury discharge rate from 113.7 to 134.2 per 100,000 (Table 1). A linear regression fit to these data predicted an average increase of 3.4 per 100,000 discharges per year ($R^2 = .93$). The largest percent increase over the eight-year period was noted among the most severe (ICD/AIS 5) injuries which nearly tripled from 1.7 to 5.0 per 100,000.

Discharge rates for young children ages 0–4 actually declined (180 per 100,000 in 1979 vs 140/100,000 in 1986) whereas rates for adolescents and young adults (ages 15–24) increased (167 vs 236), as did rates among the elderly age 75+ (145 vs 188) (Figure 1). Very little change in discharge rates was noted for adults aged 40–64. The relative risk of age-adjusted head injury discharge rates for males versus females remained relatively constant during the study period with males twice as likely as females to be hospitalized for treatment of a head injury (Table 2). The relative risk for Whites versus others, on the other hand, has declined slightly (1.3 to 1.0).

Hospital case fatality rates have declined between 1979 and 1986 (Table 3) for all severity groups with the greatest relative decrease observed among the moderately severe (ICD/AIS 3) injuries (8 to 5 percent). The percent discharged home has decreased for each severity group while the percent discharged to an extended care facility has increased. This shift in discharge disposition was most notable among individuals with the most severe injuries. Among the ICD/AIS 5 injuries, there was an 8 percent decrease in discharges to home (31 to 23 percent) and a 5 percent decrease in fatality (53 to 48 percent) accompanied by a 15 percent increase (9 to 24 percent) in discharges to an extended care facility.

Discussion

Hospital discharge abstract data generally provide a good source of information on the incidence of injury severe enough to require hospitalization,[9,11,12] and provides readily available population-based data specific to severity. Further, since these databases are maintained over time, they provide the basis for examining temporal trends in incidence.

Since the majority (80 percent) of head injuries in young children were of minor severity, the decrease may reflect changes in admission practices; moreover, in 1983, Children's Hospital in Washington, DC was designated a regional referral center for pediatric trauma, and an increasing proportion of pediatric head injuries occurring in the DC Met-

TABLE 2—Age-Adjusted Head Injury Discharge Rates by Sex and Race (per 100,000), Maryland, 1979–86

	Sex			Race		
Year	Males	Females	Relative Risk	White	Other	Relative Risk
1979	132.4	63.5	2.1	121.0	94.5	1.3
1980	131.5	61.2	2.2	119.4	91.1	1.3
1981	135.5	63.6	2.1	125.2	99.3	1.3
1982	138.7	67.6	2.1	126.5	102.9	1.2
1983	141.2	67.1	2.1	126.3	111.0	1.1
1984	144.6	71.5	2.0	134.7	116.8	1.2
1985	147.0	68.1	2.2	135.4	132.3	1.0
1986	144.2	67.4	2.1	134.0	137.7	1.0

TABLE 3—Discharge Status by Severity of Head Injury, Maryland, 1986 vs 1979

	Severity of Head Injury							
	Minor		Moderate		Serious		Severe	
Discharge Status	1979	1986	1979	1986	1979	1986	1979	1986
% Home	96	94	83	80	70	66	31	23
% Transferred to another acute care facility	1	2	4	4	5	7	7	5
% Transferred to a rehabilitation or extended care facility	1	3	5	10	7	14	9	24
% Died	2	1	8	5	18	13	53	48

ropolitan Area may have been triaged there and would not be recorded in the Maryland hospital discharge abstract database. A child restraint law enacted, in 1984 may also have contributed to the decline.[13,14]

The increased use of ICD-9CM coded discharge diagnoses as a basis for hospital payment has been shown to result in more attentive coding which might explain a shift toward more severe injuries among adolescents, young adults, and the elderly.[15] However, Maryland hospitals have been paid under a per-case payment system since the mid 1970s so that changes in coding practices are not likely to have occurred over the study period. The increased availability and use of CT (computed tomography) scans in the early 1980s may account for increased specificity in diagnosis of specific brain lesions which could result in a shift toward higher ICD/AIS scores. However, neither coding practices nor improved diagnostic procedures would likely account for the increase in overall discharge rates.

The increase in discharge rates accompanied by a decline in the hospital care fatality rates may reflect improved emergency medical services, or increased use of seat belts.[16] The increase in hospital discharge rates over the last eight years accompanied by a small decrease in the case fatality rate has resulted in an increase in the number of survivors with head injury severe enough to result in hospitalization. Several studies have documented the nature of sequelae resulting from severe as well as minor head injury.[17–25] As the number of head injury survivors increases, more hospital-based as well as community services aimed at rehabilitating the head-injured individual will be needed.

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The Impact of Legalized Abortion on Adolescent Childbearing in New York City

THEODORE J. JOYCE, PHD, AND NACI H. MOCAN, PHD

Abstract: In this paper we estimate the impact on adolescent childbearing of the liberalization of the New York State abortion law in 1970. Using Box-Jenkins time series techniques to analyze monthly data on the number of births to White and Black adolescents from January 1963 to December 1987, we found that the level of births to Black adolescents living in New York City fell 18.7 percent, approximately 142 fewer births per month, after the law became effective; the level of White births fell 14.1 percent, approximately 111 fewer births per month. Projections based on the fitted model suggest that a ban on legalized abortion today would have a major impact on adolescent childbearing in New York City as well as other parts of the country, although the magnitude of the change would vary according to local conditions. (*Am J Public Health* 1990; 80:273–278.)

Introduction

There has been great speculation in the popular press whether the US Supreme Court will overturn the 1973 decision in *Roe versus Wade*, the case which legalized abortion across the United States. A likely outcome of such a reversal is that the authority to regulate abortion will be given to individual states. To understand how a dramatic shift in the availability of legalized abortion might affect adolescent childbearing, we examined the changes in the number of births to New York City teenagers following the 1970 New York State law which liberalized abortion. It is our contention that the decline in teenage childbearing between 1970 and 1971 is a good approximation, in reverse, to what would occur today if legalized abortion were no longer available.

A number of studies have noted the decline in births after the New York State Law which liberalized abortion became effective.[1–3] However, the results in each of the studies were based on annual changes over a very short time span. Such unrefined estimates provide only a crude understanding of what might occur if the legalization of abortion were reversed. Pooled time-series, cross-sectional studies have examined the effect of liberalized abortion laws on annual, age-specific fertility rates within and across states.[4–6] The findings suggest that the legalization of abortion had an important impact on fertility rates. Again, the variation over time was limited to at most seven years.

Our study differs substantially in that it is a time-series analysis with monthly data that spans 25 years. The large number of observations allows for a more sophisticated means of fitting the data. In particular, we used Box-Jenkins methods to determine the magnitude and statistical significance of the change in adolescent childbearing that followed the liberalization of abortion in New York City in 1970. Based on the fitted model, we estimate the number of unintended births that were averted due to the widespread availability of legalized abortion. The estimates shed light on the changes in adolescent childbearing that would occur if legal abortions were to become unavailable to New York City residents.

Methods

Data

Monthly figures on the number of Black and White live births to New York City residents less than 20 years of age are from vital statistics maintained by the New York City Department of Health. Each year of individual birth records has been aggregated by month for Blacks and Whites separately. The number of birth records with unknown age in any one year was less than .03 percent; those records were deleted from the aggregation. The final series consisted of 300 monthly observations for Whites and Blacks from January 1963 to December 1987. Plots of the race-specific series are shown in Figures 1 and 2.

The analysis was limited to Whites and Blacks because ethnicity was not identified on New York City birth certificates until 1978. However, a substantial proportion of the adolescents who are White are of Hispanic origin or descent. Data from the 1970 Census indicate that 13.4 percent of all women 15 to 19 years of age were of Puerto Rican descent.[7] The number of other Hispanic adolescents in 1970 is not known. By 1980 17.0 percent of adolescents 15 to 19 years of age were Puerto Ricans, while the total proportion of all Hispanic adolescents of the same age stood at 25.9 percent. Moreover, in 1984, the first year in which data on births to Hispanic women were published, 75 percent of the White adolescents who gave birth were of Hispanic origin.[8]

The number of births as opposed to birth rates were analyzed because monthly population figures for New York City were unavailable over the period under study. Census data could have been used to estimate monthly population figures between the census years, but such crude estimates would have introduced measurement error. Furthermore, month-to-month changes in the population are minor compared to the 10 to 20 percent drop in the number of adolescent births that were observed between 1970 and 1971.

Statistical Analysis

Two approaches were used to measure the impact of the 1970 legislation on adolescent childbearing. Both approaches used Box-Jenkins time-series analysis.[9] The first approach fitted an autoregressive integrated moving average (ARIMA) model to the pre-legislation data (January 1963–June 1970). Based on the fitted model, we projected the number of adolescent births that would have been expected after July 1970 had abortion not become legal. The difference between the projected births and the actual births is one estimate of the number of births that were averted by the liberalization of abortion.

Address reprint requests to Theodore J. Joyce, PhD, National Bureau of Economic Research, 269 Mercer Street, 8th floor, New York, NY 10003. Dr. Joyce is also with the Baruch College/Mount Sinai School of Medicine, City University of New York. Dr. Mocan is with the National Bureau of Economic Research and the Manhattan College. This paper, submitted to the Journal May 25, 1989, was revised and accepted for publication August 8, 1989.

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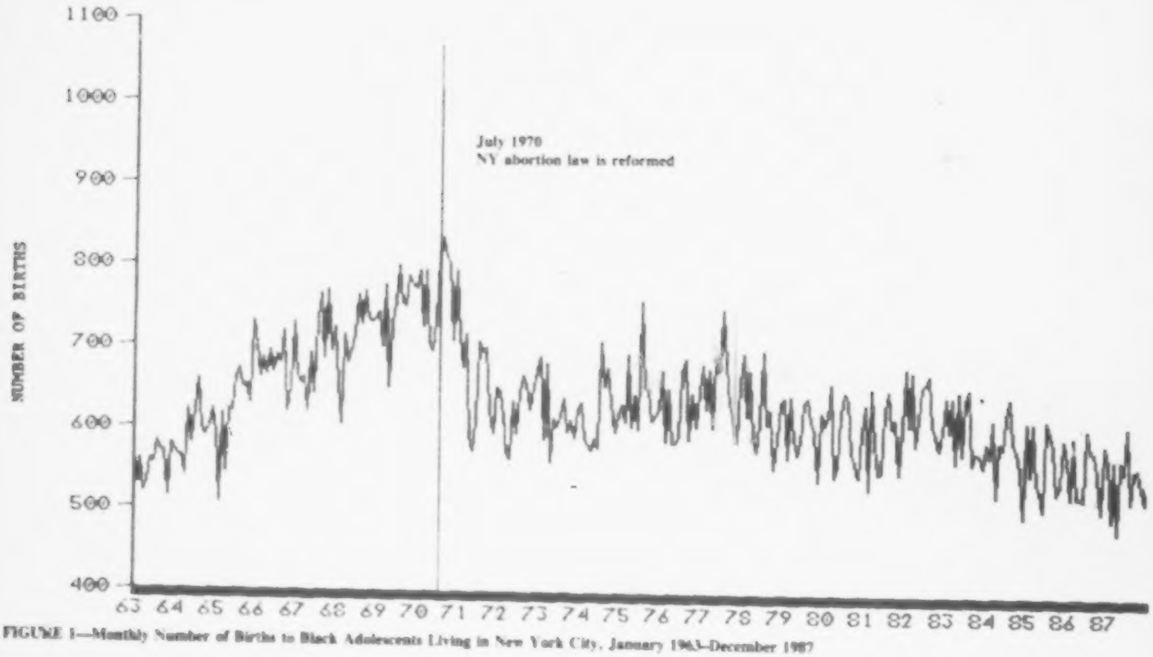

FIGURE 1—Monthly Number of Births to Black Adolescents Living in New York City, January 1963–December 1987

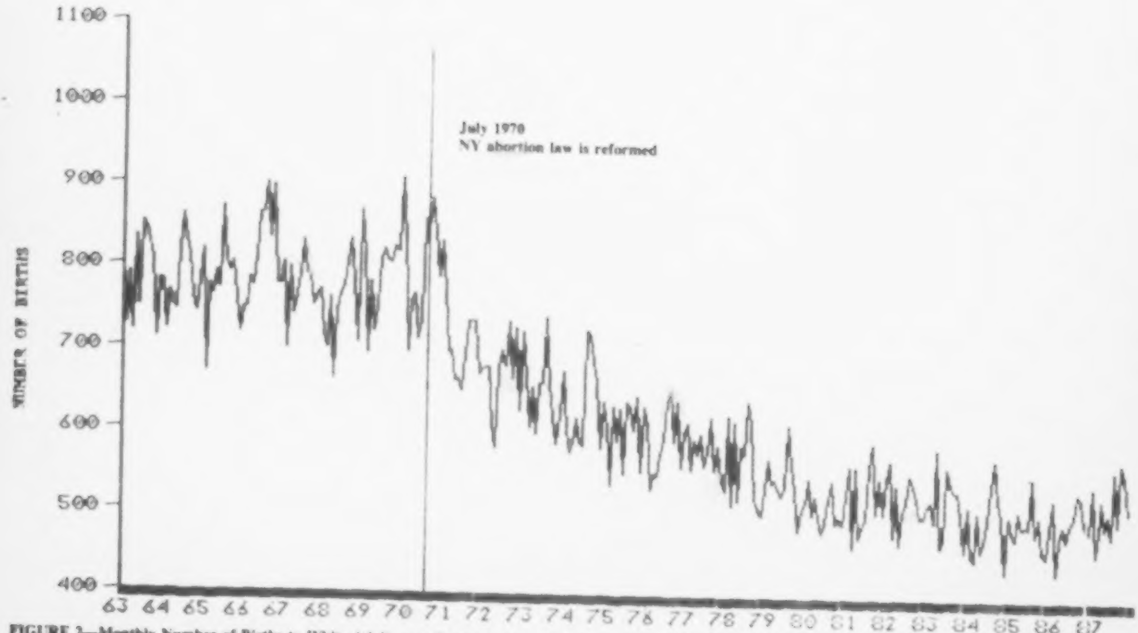

FIGURE 2—Monthly Number of Births to White Adolescents Living in New York City, January 1963–December 1987

The second approach used intervention analysis, an extension of the Box-Jenkins methodology, with the entire data set (January 1963–December 1987).[10,11] The coefficient of the intervention component measures the change in the monthly level of births between the pre- and post-intervention series. The advantage of the intervention analysis is that it is a straightforward means of determining whether the change in the series was statistically significant. Another advantage is that the magnitude of the change is based on a fitted model from observed data over the entire range of the series. The first approach is based on projections generated by the pre-intervention data only. The advantage of the first approach is that if the data were trending prior to the intervention, then the projections would incorporate the trend. The implications of the differences between the two approaches are discussed below.

To successfully apply intervention analysis, it is necessary to know the starting point of the event as well as the general shape of the response of the series to the event. The hypothesis maintained in this study is that the 1970 New York State Law which liberalized abortion had an important impact on the number of adolescent pregnancies that resulted in live births. The law became effective on July 1, 1970. However, because the law did not apply to pregnancies greater than 24 weeks, the effect of the law on the number of live births would not be observed for at least 16 to 20 weeks later. Thus, November 1970 became the starting point of the intervention. Furthermore, the full impact of the law on births to adolescents would not be realized until April 1971 when the pregnancies of the first cohort of adolescents who conceived on or after July 1, 1970 reached term. Consequently, the intervention variable was specified in such a manner that the law's impact on the number of adolescent births increased gradually from November 1970 through April 1971.

The rate at which the law's impact grew between November and April was based on the distribution of abortions by gestational age to New York City residents the first year the law was in effect. For example, 6.1 percent of all abortions to New York City residents were to women whose pregnancies were beyond the twentieth week.[2] Assuming the distribution of abortion by gestational age was the same for every month in the first year,the proportion of the law's full impact that would be felt in November was .061. Thirty percent of all abortions performed in the first year were to women whose pregnancies were between 13 and 20 weeks gestation. We assumed that 15 percent were performed between 13 and 16 weeks gestation, and the other 15 percent were performed between 17 and 20 weeks gestation. Thus, in December, the proportion of the law's full impact would be .216. This accounts for the 6.1 percent of the women who aborted in August 1970 whose pregnancies were greater than 20 weeks gestation and for the 15 percent of the abortions in July 1970 to women whose pregnancies were between 17 and 20 weeks gestation. Following this algorithm and noting that 64 percent of all abortions were performed in the first trimester, the figures for the remaining months were as follows: .361 in January, .574 in February, .785 in March, and 1.0 in April and all months thereafter.

Results

Figures 1 and 2 present the monthly number of births to Black and White New York City adolescents from January 1963 through December 1987. For Blacks, the reversal of a seven-year upward trend between 1970 and 1971 is dramatic. In the case of Whites, a relatively stationary series up to 1970 falls substantially between 1970 and 1971 and then continues downward until approximately 1986. Both figures suggest a major alteration in adolescent childbearing that is coincident with New York State's liberalized abortion law which became effective July 1970.

The ARIMA specification for the pre-intervention series is presented in Table 1. The data are expressed as natural logarithms in order to control for non-stationarity in the variance. A first-order difference transformation was applied to the logarithms of births in order to remove any trend; a twelfth-order difference was used to eliminate seasonality. Based on the autocorrelation and partial autocorrelation functions of the transformed series, Black and White births can be characterized as a first-order moving average with a first-order seasonal moving average. The coefficients of the models are displayed in Table 1. The Q-statistics in Table 1 indicate that the residuals from the estimated models are "white noise" (random variation) processes.[12] Another approach for determining if the errors are "white noise" is to test whether the first difference of the residuals follows a first-order moving-average process with the moving-average parameter equal to 1, and the first autocorrelation equal to −.5. For both Blacks and Whites this was confirmed (results not shown).

Based on the fitted model in Table 1, we projected the number of race-specific adolescent births 24 months beyond June 1970. The projections assume that abortion was not liberalized. A comparison of the actual number of births to the projected number of births is presented in Figures 3 and 4. Subtracting the actual number of births from the projected number and summing over the 24 months indicate that 4,091 Black births and 3,128 White births were averted by the legalization of abortion over the initial two-year period under the law. Note these figures include the months immediately following the new law in which the change in adolescent childbearing was relatively minor. If one estimates the *monthly* number of averted births based on the first year over which the liberalization realized its full impact (April 1971 through March 1972) then 226 Black births and 178 White births were averted per month.

To more formally test whether the 1970 law liberalizing abortion may have caused the precipitous drop in adolescent births, we used all the data to re-estimate the ARIMA structure of each series. As outlined in the previous section, an additional variable was added to the specification to

TABLE 1—Estimated ARIMA Equations for NYC Births to Black and White Adolescents 1964–1970

BLACKS	θ_1	Θ_1	$Q^a(24)$	R^2
$B_t = -\theta_1 e_{t-1} - \Theta_1 e_{t-12} + e_t$	.735 (9.16)	.801 (9.24)	15.01 .92	.63
WHITES	θ_1	Θ_1	$Q(24)$	R^2
$W_t = -\theta_1 e_{t-1} - \Theta_1 e_{t-12} + e_t$	.772 (9.38)	.644 (5.70)	19.41 .73	.56

B_t and W_t are the natural logarithms of the Black and White births respectively. θ_1 and Θ_1 are the coefficients and e_{t-1} is the error term. The numbers in parentheses are the t-ratios.

[a] The Ljung-Box Q statistic determines the randomness in autocorrelations of residual errors, and has a Chi-square distribution.[12] The numbers below the Q-statistics are the marginal significance levels, i.e., the probabilities of the null hypothesis that the autocorrelations of the errors are not different from zero.

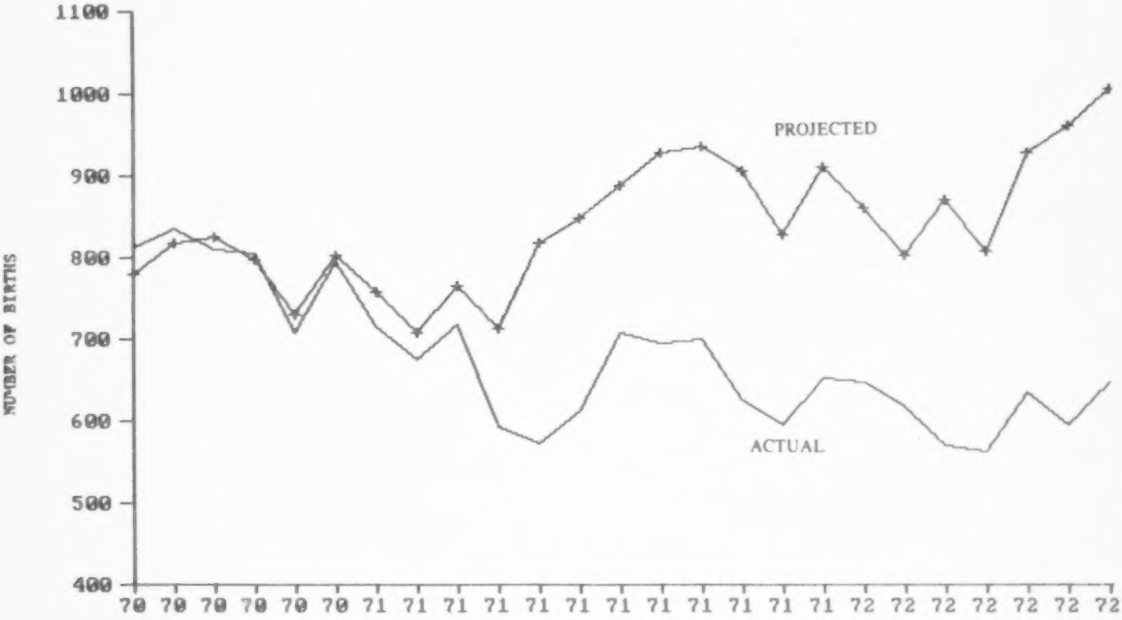

FIGURE 3—Monthly Number of Actual and Projected Births to Black Adolescents Living in New York City, July 1970–July 1972

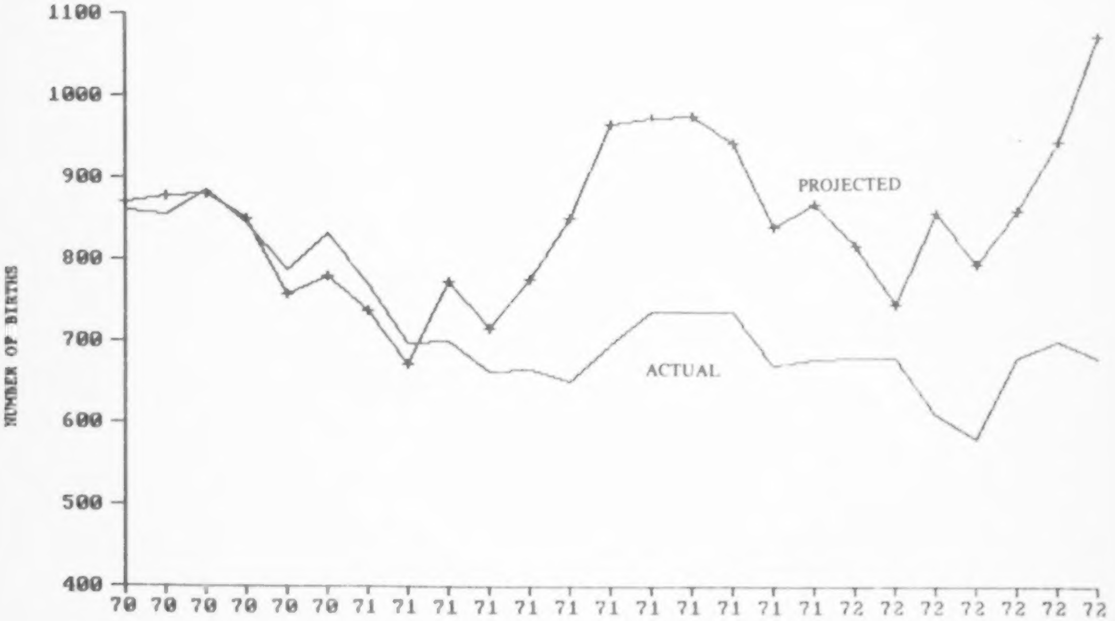

FIGURE 4—Monthly Number of Actual and Projected Births to White Adolescents Living in New York City, July 1970–July 1972

control for the impact of the law. The results are shown in Table 2. Except for the intervention component, the ARIMA structure is unchanged for Blacks. In the case of Whites, a second-order seasonal moving-average component improved the model's fit. The coefficient of the intervention variable, α, is statistically significant for Blacks and Whites. Thus, the data reveal that the decline in the level of births after October 1970 was a change that could not be explained by the normal variation in the series.

The magnitude of the change from the pre- to post-intervention level of the series can be obtained by exponentiating the coefficient on the intervention variable, α, and subtracting it from one. Expressed as a percentage, the level of Black adolescent births fell 18.7 percent after the liberalization of abortion in July 1970 (95% confidence interval = 25–12). White births fell 14.0 percent (95% CI = 20–8). If we use the average number of births over the 12 months prior to July 1970 as an estimate of the pre-intervention level of births, then 2,588 Black adolescent births and 1,998 White adolescent births were averted in the 24 months after July 1970. If the number of averted births is estimated over the first 12 months in which the law's full impact was realized (April 1971–March 1972), then 142 Black births and 111 White births were averted per month.

Discussion

Using monthly data on the number of White and Black births to New York City adolescents, we found that the liberalization of the New York State abortion law in 1970 had a substantial impact on adolescent childbearing. We also found that Blacks were more affected than Whites by the liberalization of abortion. The finding suggests that Whites may have had more access to illegal abortions than Blacks prior to the liberalization.

We used two approaches to estimate the number of births that were averted by the widespread availability of legalized abortion. We consider the estimates obtained from the intervention analysis to be conservative estimates for they do not take into account the upward trend in the number of births especially among Blacks prior to the 1970 liberalization (see Figure 1). The advantage of these estimates is that the level of births from which the changes are estimated is based on actual, as opposed to projected, data. The estimated number of averted births that were obtained by subtracting the actual number of births from the projected number of births should be viewed with some caution. Although they incorporate underlying trends, the projections are based on only seven years of data.

Finally, it should be noted that the decline in births to adolescents in the two years after July 1970 would not necessarily lead to a decline in completed fertility if the teenagers having abortions simply delayed their childbearing until a later age. However, the evidence to date indicates that early childbearers have more children and more unwanted children than women who delay childbearing.[13]

The analysis was limited to New York City because of the special circumstances surrounding the liberalization of New York abortion laws in 1970. Prior to the change in 1970, pregnant adolescents had essentially no access to legal abortion. Although Hawaii, Alaska, and Washington State has laws similar to New York's by the end of 1970, each state had residency requirements. Consequently, the magnitude of the change in adolescent childbearing among New York City residents after the 1970 law became effective was not diminished by migration to other states. The same was not true for residents of other states after the passage of the New York State law. Between July 1970 and June 1971, 75.4 percent of the 33,964 abortions performed on adolescents in New York City were to out-of-state residents.[2]

To gain some insight into the impact on teenage childbearing if legal abortion were unavailable to New York City residents, we applied the coefficients on the intervention variable in Table 2 to the projected number of births in 1988 and 1989. We adjusted for the gradual impact of a ban by the gestational age distribution of abortions prevailing today. If legal abortion were banned January 1, 1988 there would have been 2,618 (18.7 percent) additional Black births and 1,223 (14.0 percent) White births to New York City adolescents in 1988 and 1989 above what would have been expected had the laws regarding abortion remained unchanged.*

The projections are based on a number of assumptions that should be made explicit. First, we assume that the reversal of *Roe versus Wade* would result in a nationwide ban of legalized abortion. Although such an outcome is unlikely, the projections remain instructive because they provide upper-bound estimates based on the most restrictive scenario possible. This scenario cannot be dismissed lightly given that President George Bush, US Attorney General Richard Thornburgh, and Health and Human Services Secretary Louis Sullivan have publicly stated their opposition to *Roe versus Wade*.

A more likely scenario if *Roe versus Wade* were overturned is that New York State will continue to permit legalized abortion. New York is one of only 13 states plus the District of Columbia that currently funds abortions to Medicaid-eligible women and one of only eight states that does so voluntarily.[14] Nevertheless, New York's 1970 law passed the State Senate by five votes and the State Assembly by only one. The debate in both houses was tumultuous.[15]

A second assumption upon which the projections are based is that the use of abortion by adolescents in the early 1970s is unchanged. In fact, the percentage of pregnancies (live births plus induced abortions) terminated by induced abortion has risen from 40.2 percent for Whites and 47.1

TABLE 2—Estimated ARIMA Equations for NYC Births to Black and White Adolescents with the Intervention Component 1964–1987

BLACKS $B_t = -\theta_1 e_{t-1} - \Theta_1 e_{t-12} + e_t + \alpha I_t$

θ_1	Θ_1	α	Q*(48)	R^2
.776 (20.68)	.838 (24.87)	−.207 (−5.32)	34.20 .93	.64

WHITES $W_t = -\theta_1 e_{t-1} - \Theta_1 e_{t-12} - \Theta_2 e_{t-24} + e_t + \alpha I_t$

θ_1	Θ_1	Θ_2	α	Q(48)	R^2
.824 (23.94)	.740 (12.04)	.122 (1.95)	−.150 (−4.10)	42.16 .71	.61

B_t and W_t are the natural logarithms of the Black and White births respectively. θ_1 and Θ_1 are the coefficients and e_{t-1} is the error term. I_t is the dummy variable for the intervention; α is its coefficient. The numbers in the parentheses are the t-ratios.

*The Ljung-Box Q statistic determines the randomness in autocorrelations of residual errors, and has a Chi-square distribution.[12] The numbers below the Q-statistics are the marginal significance levels; i.e., the probabilities of the null hypothesis that the autocorrelations of the errors are not different from zero.

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*A detailed description of how the projections were obtained is reported in the National Bureau of Economic Research Working Paper No. 3002 by the authors.

percent for Blacks in 1972 to 54.8 and 59.2 percent, respectively, in 1986.** In this respect, our projections may underestimate the impact of a ban on legalized abortion.

One explanation for the increased use of abortion by adolescents is that the availability of abortion has engendered less effective contraceptive behavior; the more extreme version is that abortion has served as an alternative method of fertility control. Thus, if abortion were banned, the pregnancy rate may fall because of better contraceptive behavior. There is no evidence to support either explanation. A 1982 survey reported that only 0.4 percent of women at risk of an unintended pregnancy but using no contraception admitted doing so because they relied on abortion.[16] And although little is known about the relationship between abortion availability and the effective use of contraceptives, a teenager who aborted a pregnancy was less likely to become pregnant again over the next 24 months than was a comparable adolescent who carried her first pregnancy to term.[17]

As noted above, the proportion of births to White adolescents of Hispanic descent is substantial. Based on the little data that exists, Hispanic adolescents are more sexually active than their White counterparts who are not Hispanic. At the same time, Hispanic adolescents are more likely to be married and less likely to abort.[18,19] To the extent these tendencies offset each other, the proportion of Hispanics adolescents at risk of an unintended pregnancy may be similar to that of other Whites.

There is a voluminous literature on the social and economic consequences of adolescent childbearing. As the most recent and comprehensive review makes clear, adolescents who become parents will complete less schooling, have lower wages, experience greater marital instability, and be more dependent on welfare programs than their adolescent peers who delay childbearing.[13,18] Moreover, the children of teenage mothers will experience greater health, cognitive, and socioemotional difficulties.

Nor would a ban on legalized abortion be costless to taxpayers. In 1986, 64.6 percent of all adolescent births in New York City were funded by Medicaid.[20] Assuming Medicaid eligibility is a good proxy for AFDC (aid to families with dependent children) eligibility, then applying this proportion to the number of unintended births reported above indicates that 790 White and 1,691 Black teenage mothers and their children will receive AFDC in the two years after the ban. Based on the methodology described by Burt,[21] the present discounted cost in terms of Medicaid, AFDC, and food stamps of supporting a family headed by a teenager over a 20 year period is $5,560 in 1985 dollars above what it would have cost to support the same teenager and her family had she delayed childbearing until after she was 20 years of age.[21] Thus, the total marginal costs of supporting the 790 White and 1,691 Black births described above would be $13.8 million.

The percentage changes in adolescent childbearing among Whites and Blacks following a prohibition of legal abortion are clearly speculative and cannot be applied to other parts of the country because the projections are based on a set of circumstances specific to New York City in 1970. However, the number of unintended pregnancies among US adolescents strongly suggests that areas in which legal abortion is prohibited will experience substantial increases in the number of births to teenagers. The magnitude of the change will vary by area because of differences in the use of abortion prior to a ban, the proximity to areas where abortion remains legal, and the availability of illegal abortions.

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**The figures are based on data provided by the New York City Department of Health.

An Evaluation of External Cause-of-Injury Codes Using Hospital Records from the Indian Health Service, 1985

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Abstract: To evaluate the usefulness of International Classification of Diseases external cause-of-injury and poisoning codes (E codes) for public health surveillance of nonfatal injuries, we analyzed E codes from Indian Health Service (IHS) hospital records. E codes for unknown or unspecified causes were used for 25 percent of records. At two hospitals, 63 percent of E codes assigned by independent coders agreed; another 18 percent matched on general cause-of-injury groups. With uniform guidelines and increased training, E coding could provide a valuable, cost-effective method of quantifying and characterizing severe, nonfatal injuries. (*Am J Public Health* 1990; 80:279–281.)

Introduction

Because external causes of injury and poisoning are routinely coded on death certificates under the International Classification of Diseases (ICD) external cause-of-injury and poisoning coding scheme (E codes),[1] we have learned a great deal about fatal injuries. We know much less, however, about the causes of the many more numerous nonfatal injuries, partially because we lack the appropriate data. Although hospital discharge records provide a potential mechanism for public health surveillance of nonfatal injuries severe enough to require hospitalization, most hospitals do not routinely E code discharge data. Recently, influential health policy and advisory groups, including the National Academy of Sciences and the Council of State and Territorial Epidemiologists, have advocated routine E coding of hospital discharge data to link injury data collection and the development and implementation of intervention strategies.[2–4]

Presently, population-based E coded hospital discharge data are already available in the Indian Health Service (IHS). To examine a practical application of hospital discharge data E coding as well as the strengths and limitations of using E codes for injury surveillance, we analyzed IHS records to determine the level and quality of E coding, the reliability of E codes assigned by different coders, and the feasibility of using E coded hospital discharge data as an injury surveillance tool.

Methods

Discharge records for eligible American Indians and Alaska Natives hospitalized at IHS facilities from October 1, 1984—September 30, 1985, were selected for analysis. We excluded records for medical care provided by outside contractors and at tribal facilities (27.9 percent of admissions[5]) because they are not reliably E coded.

Medical records for all hospitalizations were coded by on-site medical records personnel at each of 46 IHS hospitals using the standard ICD (ninth revision)-Clinical Modification (ICD-9-CM)[6] codes and rules. Records were coded after discharge based on clinical assessment, diagnostic and therapeutic procedures, and other health services information collected during the hospitalization and recorded on the chart. Each IHS record was coded for: 1) up to six clinical codes; and 2) one cause-of-injury code, if the hospitalization was injury-related.

The ICD hospital indexing system includes the Classification of Diseases and Injuries (often called "N codes"); two supplementary classifications: Factors Influencing Health Status and Contact with Health Service (V codes*) and External Causes of Injury and Poisoning (E codes); and four appendices (Morphology of Neoplasms, Glossary of Mental Disorders, Classification of Drugs, and Classification of Industrial Accidents). Diagnostic and therapeutic procedures are coded with yet another classification scheme, the Classification of Procedures in Medicine.

The cause-of-injury/poisoning E code has several components. All E codes have three digits following an "E." Three-digit E codes can be grouped according to general external cause or intent (e.g., E880-E888: accidental** falls, E950–E959: suicide and self-inflicted injury). Most three-digit E codes also require a fourth digit subdivision (EXXX.0–EXXX.9) to identify the role of the injured person or to provide additional cause-specific information.

For this analysis, we selected injury records by two independent methods: 1) records with injury and poisoning diagnosis and health services codes (ICD-9-CM codes 800–999 and V71.3–V71.6); and 2) records receiving cause-of-injury E codes (E800–E999). We then compared the files to measure the level of injury coding. To evaluate the reliability of IHS E codes, we reviewed E coded records for two hospitals, abstracted cause-of-injury information, and recoded the external causes of injury and poisoning. The author (SMS) who recoded records was blinded to the original codes.

Results

Of the 10,921 (13.9 percent) IHS hospital discharge records with at least one injury-related diagnosis or health services code, 99.3 percent were E coded. All records with an injury diagnosis code (800–999) had an associated E code;

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*V codes are used for occasions when a person encounters the health care system for a specific purpose other than illness or injury (e.g., immunization or pregnancy). For injuries, certain V codes may be appropriate for persons hospitalized for observation following a traumatic event.

**Although use of the term "accidental" denotes an unpreventable event and, therefore, is to be avoided, unintentional injuries are labeled "accidental" in the ICD.

52 of 128 records selected by an injury health services codes (V71.3–V71.6) were E coded.

Of the 11,063 E coded records selected independent of diagnosis codes, 10,845 (98.0 percent) had at least one injury or poisoning diagnosis or health services code. For the other 218 E-coded records, the most frequent conditions requiring hospitalization were adverse effects of medical/surgical care (N = 106) and poisoning (N = 9). These 11,063 hospitalizations were required for 9,555 individuals.

More than 75 percent of all injury hospitalizations were in six broad external cause groups: transport-related incidents, falls, acts of violence, adverse effects of medical and surgical care, other accidents (being struck, caught, or cut), and suicides/attempts (Table 1). Transport-related events and falls together accounted for more than a third of all injury hospitalizations.

About 25 percent of IHS injury hospitalization records were coded using three-digit E codes for general or unspecified causes of injury, reflecting a lack of more detailed information about the injury events. Most frequently, records for falls, transport-related injuries, and assaults could be coded with only general E codes. For 540 records, the underlying cause of injury was not available, and coders used two non-specific codes, E887 (fracture, cause unspecified) and E928 (other and unspecified environmental and accidental causes).

Fourth-digit codes were assigned where appropriate. Coders assigned a fourth-digit code to all three-digits E codes requiring the additional character. For those three-digit E codes without an associated fourth-digit, IHS appropriately did not assign a fourth digit. For 22.6 percent of records with E codes requiring a fourth digit, a nonspecific fourth digit code was assigned because sufficient information to code the fourth digit was not provided in the record.

The E coding reliability evaluation was based on injury hospitalization records from two IHS Service Units. Ninety-seven percent of 333 discharge records receiving E codes were available to abstract and recode. We assessed coding agreement on a hierarchical scale. When we compared recoded E codes with those assigned by IHS medical records personnel, 59 percent agreed completely (agreement to the three-digit level and agreement to the four-digit level where appropriate). Comparison of the three-digit agreement hierarchy is shown in Table 2. Overall, in 63 percent of the records, the original and recoded three-digit E codes matched. Eighty-one percent of records had matching three-digit codes or similar codes from an external cause group. An additional 5 percent of records had original and recoded E codes that fell under the same ICD-9-CM rubric of intent (i.e., unintentional, intentional, or unknown intent of injury). Neither three-digit E code, external cause group, nor intent agreed for 14 percent of records. For most of the nonmatching sets of codes, medical care providers recorded different scenarios of the injury event on an individual chart or E codes were selected on the basis of recorded injury event circumstances that did not fit well within the ICD-9-CM classification scheme (e.g., Should a fall out of an automobile be coded as a fall or as a transport-related event?).

TABLE 1—Frequency of Three-Digit External Cause-of-Injury Codes from Hospitalization Records,* by Injury Group, Indian Health Service, Fiscal Year 1985

E Codes	External Cause Group	Frequency	Percent	Cumulative Percent
E800–E848	Transport-related	1896	17.1	17.1
E880–E888	Falls	1871	16.9	34.1
E960–E978,E990–E999	Assaults/legal/war	1606	14.5	48.6
E870–E879,E930–E949	Adverse effects of care	1367	12.4	60.9
E916–E920	Struck/caught/cut	973	8.8	69.7
E950–E959	Suicides/attempts	805	7.3	77.0
E980–E989	Intent unknown	477	4.3	81.3
E900–E909	Environmental	456	4.1	85.4
E928–E929	Other/late effects	455	4.1	89.5
E850–E869	Poisonings	366	3.3	92.9
E927	Overexertion	221	2.0	94.8
All Other E Codes		570	5.2	100.0

*N = 11,063

TABLE 2—Percent Agreement of External Cause-of-Injury Codes for Injury Hospitalization Records, Indian Health Service

	Percent Agreement*		
Level of Agreement	Service Unit A (N = 240)	Service Unit B (N = 83)	Cumulative Percent (N = 323)
Three-digit E code	60	71	63
External cause group	76	91	81
Intent**	83	92	86

*A mutually exclusive hierarchy. For each hospital record, the E code assigned by IHS medical records personnel was compared to the E code assigned by one of the authors. Codes matching on the external cause group are those for which both coders assigned three-digit E codes from the same group of external causes (e.g., transport-related incident [E800–E848], falls [E880–E888], or assaults [E960–E978]) but did not assign identical three-digit E codes.

**Codes matching on intent are those for which both coders assigned E codes of the same intent classification (i.e., unintentional, intentional, or unknown intent) but did not assign identical three-digit E codes or codes from the same external cause group.

Discussion

Because proposals to require universal E coding of hospital discharge data are gaining prominence,[2–4,7,8] this examination of E code use in the IHS is valuable for public health practitioners and other researchers interested in injury surveillance and data collection. Few other opportunities exist for us to test the feasibility of requiring universal E coding and to understand the strengths and limitations of injury cause information obtained from E coded records. As shown in this evaluation, routine E coding of IHS hospital discharge data was a valuable, cost-effective method of quantifying and characterizing severe, nonfatal injuries. E codes provided us with a clear distribution of the underlying causes of injuries requiring hospitalization in this population, thus permitting us to identify priority injury problems. Community-based injury prevention programs, based in part on external cause information extracted from hospital records, are currently being coordinated and conducted by trained injury prevention professionals throughout IHS.[9] The E coded records are also valuable for evaluating the effectiveness of these prevention programs.

E codes were reproducible. Although the specific three- and four-digit codes assigned by independent coders sometimes differed, codes agreed closely according to the external cause group. The extent of E coding agreement compares favorably with previous evaluations of the reliability of death certificate and cancer registry information. In Connecticut, an analysis of autopsy reports and corresponding death certificates found major disagreement leading to reclassification of the underlying cause of death for 29 percent of deaths.[10] For an additional 26 percent of deaths, the death certificate and autopsy report agreed on the major disease category but attributed the death to a different specific disease. As part of the Third National Cancer Survey, death certificates for 48,826 cases of single primary cancer were compared to hospital diagnoses.[11] The underlying cause of death was found to be accurate in 65 percent of cases. In a comparison of data collected from interviews and those reported to the Missouri cancer registry,[12] 70 percent of occupation, 72 percent of industry, 83 percent of smoking status, and 65 percent of alcohol consumption codes agreed. In spite of these limitations, coded data from death certificates and cancer registries have been used extensively in public health practice. E coded hospital discharge data can be of similar value.

With increased attention to the value of E codes, reliability can be improved. E code users and potential users would benefit from standardized E coding guidelines and definitions, increased availability of E code training, and more rigorous evaluation of the current E code scheme. For example, because only 40.6 percent of hospitalizations for observation following trauma (V71.3–V71.6) received E codes in this analysis, we assume that the use of E codes for records coded with health services codes (V codes), and perhaps other injury codes, was not consistent. Due to reimbursement, hospitals have financial incentives to provide adequate training for coders to ensure accurate diagnosis and procedure coding. As more states require E coding of uniform hospital discharge data,[7] the need for uniform E coding procedures and training will increase.

About a fourth of IHS injury records were coded with E codes for general or unspecified external causes. These general or unspecified codes are much less useful for public health surveillance and the development of injury prevention programs. The nonspecific E codes were used most often because more detailed cause-of-injury information was not recorded in the medical charts. Medical record keeping is, understandably, focused on recording clinical assessment and treatment. Generally, health care providers are not aware of the other ongoing or potential uses for medical record data or of the priority data items to collect for public health surveillance. By presenting studies such as this one, health care providers can be made more aware of the crucial role of the medical record in the flow of data through epidemiologic and statistical processes to health care policy development and evaluation.

Because almost all IHS discharge records with injury diagnosis codes received E codes (including multiple injury-related hospitalizations for individuals in the same fiscal year), these E coded records do not represent injury incidence. We could not determine whether the multiple admissions for individuals were due to multiple injury events or follow-up medical and surgical care for previous injuries. E codes could describe injury incidence, however, if coding rules were standardized, restricting E codes to the initial hospitalization following an injury, or if they were modified to include a separate code identifying the initial hospitalization. We could then code records for subsequent hospitalizations consistently using the existing diagnosis codes available for follow-up treatment and late effects of injuries.

Although E coded hospital discharge data have great value for injury surveillance, they have more limited value for epidemiologic research. E codes can help us meet some of the priority objectives for public health surveillance—ascertaining cases, quantifying and characterizing injuries, and monitoring trends in a defined community.[10] Because the number of E codes for certain specific injury events is restricted or absent, the detail needed for focused epidemiologic research is limited. To obtain more detailed cause-of-injury data or information about specific injury problems of interest (e.g., injuries associated with all-terrain vehicles) we will require other surveillance tools or analytic epidemiologic research.

Information about underlying causes is essential for the development, implementation, and evaluation of injury intervention strategies. E coding of hospital records, as performed in the IHS, provides much of this underlying cause information. By extending hospital record E coding to other health care systems, we can develop a standardized external cause data base for prevention programs.

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Cardiovascular Deaths among Alaskan Natives, 1980–86

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Abstract: Average annual, age-adjusted death rates from cardiovascular diseases and atherosclerosis from 1980–86 among Alaskan Natives were lower than rates among other Alaskans (162.0 vs 242.1; RR = 0.67), while death rates from other causes were higher (954.4 vs 618.6; RR = 1.54). These suggest that Alaskan Natives have less cardiovascular disease than other populations. Additional research on the role of marine omega-3 fatty acids is needed. (*Am J Public Health* 1990; 80:282–285.)

Introduction

Numerous reviews on the role of diets rich in omega-3 fatty acids and the incidence of atherosclerosis and coronary artery disease recently have been published, reflecting widespread interest and enthusiasm over the potential beneficial health effects of omega-3 fatty acid consumption.[1–13]

Careful review of published literature reveals that the epidemiologic studies that show a low incidence of atherosclerosis and cardiovascular disease in arctic peoples, particularly, rest on fragile evidence.[6–9,14,15] For example, earlier works in Alaska to investigate causes of deaths were based on a limited number of autopsies conducted by non-pathologists using no standardized protocols without histopathology, and the two studies reached opposite conclusions about incidence of atherosclerosis among Alaskan Natives.[8,9] A study of death records for 1955–65 found lower mortality rates from heart disease among Alaskan Natives than the US population.[15] This study presents additional data based on death certificates on the causes of death among Alaskan Natives.

Methods

All Alaskan death certificates for 1980 through 1986 (N = 13,534) were reviewed. The underlying cause of death and all other conditions that contributed to the death as recorded on the death certificate were coded using the *Ninth International Classification of Diseases—Ninth Revision* (ICD-9). As a validation of the accuracy of coding, for deaths occurring during 1980–82, we manually verified our codes against those furnished by the National Center for Health Statistics (NCHS). Population estimates were obtained from the Alaska Department of Labor and the US Bureau of Census.[16] Alaskan Natives, consistent with designations of the Alaska Department of Labor and the US Bureau of Census, included Eskimo (Inupiaq and Yu'pik), American Indian (Athebaskan, Tlinqit, Haida, Tsimshian and other North American tribes), and Aleut.

We compared average annual death rates between Alaskan Natives and others according to major causes of death, using age-adjusted rates calculated by the direct method using the 1980 US Census as the standard. We compared cause-specific, age-adjusted death rates using the underlying cause of death, only, and by using contributory causes. Groupings of ICD codes are shown in Table 1 and the Appendix. Statistical analyses were done using Epi Info, Version 3, Division of Surveillance and Epidemiologic Studies, Epidemiology Program Office, Centers for Disease Control, Atlanta, Georgia.

Results

Alaskan Natives had a lower age-adjusted mortality rate from atherosclerosis and cardiovascular disease compared to other Alaskans (162.0 vs 242.1; RR = 0.67). Because the differences in death rates between underlying cause and multiple cause were so similar, only data for underlying cause are presented (Table 1). The only major difference in cardiovascular death rates by sex was for the comparison of ICD code 410, acute myocardial infarction, for underlying cause only, where female Natives had a slightly higher average annual, age-adjusted death rate than other females. For all other comparisons of cardiovascular disease-related groups, using underlying cause of death only, multiple causes, male, female, or total population, Alaskan Natives had consistently about 60 percent of the death rate of others.

In contrast, the death rates from infections, respiratory diseases, and neoplasms were all higher among Alaskan Natives than in others. The death rate from cardiovascular disease rose strikingly with age for both Natives and others (Figure 1) and was consistently present among the various cardiovascular-related disease, ICD code groups, by underlying cause and multiple cause, and in both males and females.

Interestingly, there was only one Native death due to aortic aneurysm (ICD code 441) compared to 88 deaths among others (RR = 23.8). The death rate due to all cerebrovascular disease (ICD codes 430–438) was slightly higher among Natives (RR = 1.08) and were primarily related to a higher rate of deaths due to cerebrovascular hemorrhage (ICD codes 430–432; RR = 1.36).

Discussion

This analysis of death certificate data provides additional evidence that arctic indigenous people have strikingly lower death rates from cardiovascular disease than others. The death rate due to cerebrovascular hemorrhage was slightly higher among Alaskan Natives than others, and similar to the overall difference in mortality rates between Natives and others. Other investigators have reported a high rate of cerebrovascular hemorrhage among Eskimos[14,17] and Japanese[18] who consumed high quantities of fish, raising concerns that increased use of dietary consumption of fish oils may lead to the unwanted consequence of cerebral hemorrhage. Our data do not support this concern.

Limitations of death certificates are well known.[19–22] No data exist about their accuracy in Alaska or the correlation of recorded causes of death with autopsy results. Autopsy

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TABLE 1—Causes of Death in Alaskans, 1980–86

Cause of Death ICD-9 Codes	Gender	Number of Deaths		Rate*		RR**	95% Confidence Interval
		Native	Other	Native	Other		
Cardiovascular Diseases and Atherosclerosis							
410	Male	74	323	54.3	60.6	0.90	0.55,1.44
	Female	42	106	43.4	32.8	1.32	0.76,2.35
	Total	**116**	**429**	**50.2**	**49.5**	1.01	**0.71,1.46**
410–414	Male	205	1,339	154.1	245.7	0.63	0.47,082
	Female	87	448	90.8	140.0	0.65	0.45,0.94
	Total	**292**	**1,787**	**127.9**	**203.0**	0.63	**0.51,0.79**
401–404, 410–414, 429.2, 440	Male	245	1,483	186.2	279.3	0.67	0.52,0.86
	Female	118	568	125.9	182.1	0.69	0.50,0.95
	Total	**363**	**2,051**	**162.0**	**242.1**	0.67	**0.55,0.81**
Cerebrovascular Diseases							
431	Male	6	24	3.6	3.8	0.95	0.09,5.98
	Female	11	24	6.9	6.2	1.11	0.22,4.38
	Total	**17**	**48**	**5.1**	**4.9**	1.04	**0.41,3.51**
430–432	Male	14	46	8.4	5.8	1.45	0.42,5.30
	Female	24	59	16.1	12.5	1.29	0.44,3.06
	Total	**38**	**105**	**12.1**	**8.9**	1.36	**0.60,2.79**
430–438	Male	58	180	43.7	38.7	1.13	0.63,1.90
	Female	65	194	60.0	58.0	1.03	0.63,1.63
	Total	**123**	**374**	**52.0**	**48.2**	1.08	**0.75,1.53**
Aneurysm 441	Male	1	68	0.8	15.5	0.05	0.02,1.33
	Female	0	20	0	6.5	—	
	Total	**1**	**88**	**0.5**	**11.9**	0.04	**0.02,0.87**
Infections 001–139	Male	21	59	12.7	7.0	1.81	0.53,4.81
	Female	18	29	11.9	5.6	2.13	0.67,6.82
	Total	**39**	**88**	**12.5**	**6.5**	1.92	**0.92,4.19**
Respiratory 460–519	Male	139	270	97.1	61.9	1.57	1.07,2.27
	Female	100	214	83.9	61.4	1.37	0.92,2.08
	Total	**239**	**484**	**93.2**	**63.3**	1.47	**1.13,1.95**
Neoplasms 140–239	Male	300	1,051	223.6	198.4	1.13	0.89,1.44
	Female	243	823	216.3	182.5	1.19	0.92,1.53
	Total	**543**	**1,874**	**224.8**	**195.7**	1.15	**0.97,1.36**
Intentional Injuries E950-E969	Male	302	572	124.6	40.8	3.10	2.13,4.40
	Female	86	158	34.5	13.6	2.50	1.29,5.11
	Total	**388**	**730**	**79.6**	**28.5**	2.80	**2.02,3.85**
Unintentional Injuries E000-E949, E970-E999	Male	615	1,829	266.3	133.3	2.00	1.58,2.52
	Female	201	400	93.4	43.5	2.10	1.44,3.26
	Total	**816**	**2,229**	**180.3**	**92.7**	1.90	**1.59,2.38**
All except Cardiovascular	Male	1,990	5,203	1,071.6	665.4	1.64	1.46,1.83
	Female	1,173	2,684	817.4	553.0	1.48	1.30,1.69
	Total	**3,163**	**7,887**	**954.4**	**618.6**	1.54	**1.42,1.68**
All Causes	Male	2,235	6,686	1,257.8	934.7	1.35	1.21,1.49
	Female	1,291	3,252	943.3	735.1	1.28	1.14,1.45
	Total	**3,526**	**9,938**	**1,116.4**	**860.8**	1.30	**1.20,1.40**

*Average annual age-adjusted rate per 100,000 by direct method to 1980 U.S. population
**Native rate/Other rate

results, if available, were not routinely used to determine cause of death that was recorded on the death certificate by the attending physician. ICD coding was done by physician epidemiologists who were not trained nosologists but for another study and prior to involvement with work on diet and atherosclerosis—therefore, potential bias in coding cardiovascular related disease was minimized. We coded all death certificates because previous studies revealed many incorrectly coded certificates and because only the underlying cause of death had been coded. We analyzed our data using our codes; results using multiple cause coding were similar to results using underlying cause only. An analysis of NCHS codes using underlying cause only also showed similar results.

Many recent studies have focused on health effects of fish consumption.[1–5,13,23] Bang emphasized the role of Eskimo diet rich in marine mammal fatty acids, especially seal and whale, and these foods generally were eaten raw.[24] Whether the same health benefits can be anticipated from raw fish and from cooked fish is less clear. Consumption of

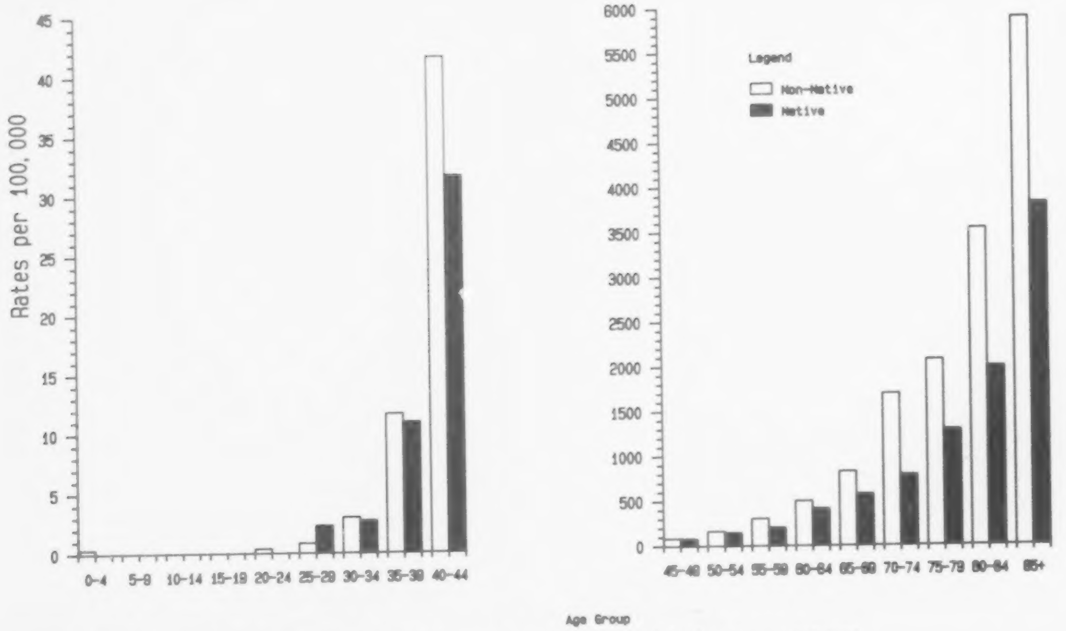

FIGURE 1—Cardiovascular Deaths in Alaska 1980–86 (average annual rate per 100,000 population by age group and race)

salmon remains high and seal, walrus, and whale consumption, both raw and cooked, remains an important part of Alaskan Natives' subsistence diet.

As early as 1900, diet changes among Alaskan Natives were documented as a result of contact with whalers and later with missionaries.[1] Since the 1950s, there has been a dramatic increase in smoking and alcohol consumption. Since the 1960s, western diets increasingly have become available. Since the 1970s, use of snuff has increased.[25] In view of these many life-style changes, it is even more remarkable to find such striking differences in cardiovascular mortality.

The evidence presented here does not establish that the cause of the difference in cardiovascular mortality rates is due to dietary intake of omega-3 fatty acids from marine mammals and fish. It is possible that the differences could be due to genetic factors, or that the observed differences are a result to the effect of competing risks[21,22] acting disproportionately on Natives, or to inaccuracies inherent in death certificate studies. For example, Alaskan Natives may have differences from other Alaskans in the nature and frequency of medical system contacts, the likelihood of receiving care from physicians of similar background and training, availability of diagnostic services, death certification practices in rural compared to urban areas, classification of race on the death certificate, and differences in accuracy of census data. Additional research is needed to pursue accumulating evidence of the beneficial effects that may exist from consumption of omega-3 fatty acids.

ACKNOWLEDGMENTS

The author thanks Eitel Dunaway for computer analysis and graphics production; Diane Ingle and Sue Anne Jenkerson, RNC, MSN, FNC, for data entry; and W. Gary Hlady, MD, for assistance in coding causes of death. I am most grateful to the Alaska Bureau of Vital Statistics whose cooperation and hospitality made this study possible.

APPENDIX

ICD Code	*Specific Diagnosis*
401	Essential hypertension
402	Hypertensive heart disease
403	Hypertensive renal disease
404	Hypertensive heart and renal disease
410	Acute myocardial infarction
411	Other acute and subacute forms of ischemic heart disease
412	Old myocardial infarction
413	Angina Pectoris
414	Other forms of chronic ischemic heart disease
429.2	Cardiovascular disease, unspecified
440	Atherosclerosis

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ERRATA

In: Wells JK, Williams AF, Fields M: Coverage gaps in seat belt use laws. Am J Public Health 1989; 79:332–333. Table 1 which appeared on page 333 last year is in error. In the listing of all the states, Utah was inadvertently omitted and the wrong data presented for Tennessee and Texas. All other states and data were correct. Printed below are the correct figures for Tennessee, Texas, and Utah. We regret the error.

TABLE 1—Correction

		Not Subsequently Covered		Total Not Covered	
State	Total Deaths	Rear Seat Occupants	Light Truck Van Occupants	N	%
Tennessee	3,013	111	0	111	4
Texas	12,564	784	0	784	6
Utah	882	71	0	71	8

* * *

In: Popkin BM, Akin JS, Flieger W, Wong EL: Breastfeeding trends in the Philippines, 1973 and 1983. Am J Public Health 1989; 79(1):32–35. The authors wish to call attention to an error in the coding manual for a data set they used in that paper. Although the error does not change the substantive results, it does change slightly the results presented in their Table 2 (page 34) under the subheading of "Mother's Education (years)," lines 4 and 5 for "7–9" years and "10± years". The corrected results show a slightly greater decline among those with 7–10 years of schooling and with more than 11 years of schooling, as shown in the two lines in the table correction below.

TABLE 2—Correction

		Incidence			95% Confidence Interval of Difference	
Characteristics	Average Sample Size	1973	1983	Unadjusted Difference*	Lower Limit	Upper Limit
7–10	763	86.7	81.2	−5.5	−9.1	−1.9
11±	1058	66.2	72.6	6.4	1.0	11.8

Salmonella arizona Infections in Latinos Associated with Rattlesnake Folk Medicine

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Abstract: In 1987 two Los Angeles County (California) hospitals reported four Latino patients with serious *Salmonella arizona* (*Salmonella* subgroup 3) infections who gave a medical history of taking rattlesnake capsules prior to illness. Capsules supplied by the patients or household members grew *Salmonella arizona*. We reviewed surveillance data for this *Salmonella* species and conducted a case-control study to determine the magnitude of this public health problem. Eighteen (82 percent) of the 22 Latino cases in 1986 and 1987 who were questioned reported ingesting snake capsules compared to two (8 percent) of 24 matched Latino controls with non-subgroup 3 salmonellosis or shigellosis (matched pair odds ratio = 18.0, CI = 4.2, 76.3). An average of 18 cases per year of *Salmonella arizona* were reported in the county between 1980 and 1987. In this investigation the majority of *S. arizona* cases reporting snake capsule ingestion had underlying illnesses such as acquired immunodeficiency syndrome (AIDS), diabetes, arthritis, cancer. The capsules were obtained primarily from Tijuana, Mexico and from Los Angeles, California pharmacies in Latino neighborhoods. Despite publicity and attempts to remove the capsules from sale in California, *Salmonella arizona* cases associated with snake-capsule ingestion continue to occur. (*Am J Public Health* 1990; 80:286–289.)

Introduction

In early 1987, two hospitals in Los Angeles County (California) reported four cases of *Salmonella arizona* (*Salmonella* subgroup 3, formerly *Arizona hinshawii*) infection in Latino patients to the Los Angeles County Department of Health Services. All patients gave a history of ingesting rattlesnake capsules prior to the onset of illness; in all four cases, culture of the rattlesnake capsules from the patients' supply or from their household grew *Salmonella arizona* and other enteric bacteria.[1] The patients had underlying chronic medical illnesses including acquired immunodeficiency syndrome (AIDS) and cancer, and the bacteria were isolated from extragastrointestinal sites (blood, pleural fluid, lymph node). We conducted an investigation to define further the epidemiology of rattlesnake-capsule ingestion and *Salmonella arizona* infections in Los Angeles County.

Methods

Surveillance data for *A. hinshawii* and *S. arizona* infections were reviewed for 1980–87. A matched case-control study was performed in late 1987 and early 1988. All Latino cases of *S. arizona* reported to the Los Angeles County Department of Health Services during 1986 and 1987 were eligible. Each case was matched to one or more persons with either *Salmonella typhimurium*, *Salmonella enteritidis*, or *Shigella*. *Salmonella* controls were chosen from 1986 and 1987 while *Shigella* controls were chosen from 1987. Cases were matched on race and age. Controls were at least 17 years old.

All cases of salmonellosis and shigellosis occurring in Los Angeles County are routinely administered a standard two-page questionnaire covering demographics, clinical information, and potential exposures. These data were supplemented by non-blinded administration of a telephone questionnaire which contained information regarding rattlesnake-pill ingestion, underlying illness, and folk medicine usage. Most interviews were conducted in Spanish by bilingual investigators. Serotype data for all *S. arizona* isolates submitted to the Los Angeles County Public Health laboratories during 1986–87 were reviewed.

Time of exposure was defined as the date to the nearest week, month, and/or year in some cases that the patient reported taking rattlesnake capsules. Since many of the patients were interviewed six months or more after their illness, they often had to estimate these dates. Estimated incubation period was defined as the interval between most recent exposure to rattlesnake pills and date of onset of illness or date of first positive culture, if illness onset date was unclear.

Unless otherwise specified, Mantel-Haenszel matched analysis was used to determine odds ratios and 95 percent confidence intervals (CI).

Results

An average of 18 cases a year of *S. arizona* were reported in Los Angeles County during 1980–87 (Figure 1). The number of cases ranged from nine in 1984 to 33 in 1987. In 1986–87, 45 (85 percent) of 53 such cases were Latino, while 208 (31 percent) of 673 cases with other common *Salmonella* serotypes were Latino (95 percent CI of proportion difference = 0.43, 0.65).

Twenty-two cases were interviewed with the supplemental questionnaire. Not surprisingly, we were able to contact more of the 1987 cases (18 of 29) than 1986 cases (four of 16). The age and geographic distributions of *S. arizona* cases not included were similar to those of interviewed cases. Matched controls included 12 *Salmonella* patients from 1986–87 and 12 *Shigella* patients from 1987. The pool of potential *Salmonella* controls in the proper age range was small. Age matching of cases and controls was ±10 years for 96 percent of the matches; the largest age difference was 14 years.

The median age of cases was 44 years with a range of 16 to 61 years. The male-to-female ratio of interviewed cases was 2.7:1 while that for all Latino *S. arizona* cases was 1.5:1. Cases lived throughout the County but were concentrated in neighborhoods within central Los Angeles. The birthplace or national background of rattlesnake capsule-exposed cases was predominantly Mexico (72 percent).

Eighteen (82 percent) of the 22 cases gave a history of rattlesnake-capsule exposure compared to two (8 percent) of

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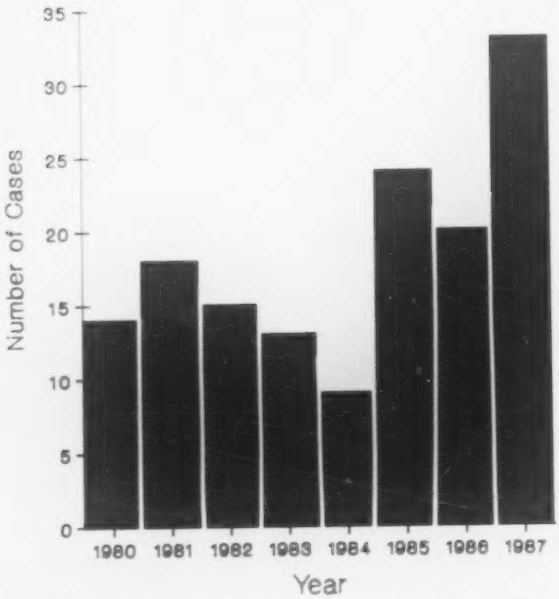

FIGURE 1—Cases of *Salmonella arizona* by Year, Los Angeles County, California 1980–87

24 matched controls (OR = 18.0, CI = 4.2, 76.3). The attributable fraction of *S. arizona* cases due to rattlesnake-capsule exposure in the Latino population is 77 percent.

Ten (56 per cent) of the 18 cases took fewer than 10 capsules over a period of a week or less. Five patients took only two or three pills. One patient ingested capsules daily for four months. The interval between exposure and onset of symptoms or first positive culture ranged from within a week to four years. Eleven (73 percent) of the 15 cases for whom an incubation period could be estimated became ill within three months. Two of the patients with longer estimated incubation periods had perirectal abscesses (Table 1). No correlation was observed between number of pills ingested and incubation period or site of infection.

The reasons patients used rattlesnake capsules included treatments for AIDS, arthritis, blood disorders, cancer, infections, itchy feet, sinus and skin conditions, and diarrhea. The majority of rattlesnake capsules were obtained from Tijuana, Mexico and from Los Angeles, California pharmacies. Patients also reported purchasing rattlesnake capsules in other Mexican border cities, in Central America, from a street vendor, at a local "swap meet" and from a friend. Capsules sold in Los Angeles cost about 20 cents per capsule. Forty-one percent of cases and 54 percent of the controls used other common folk medicines including yerba buena and other herbs and teas.

Feces, blood, and lung or pleural fluid were the most common sites of *S. arizona* isolation (Table 2). All three patients with multiple positive sites of *S. arizona* isolates had AIDS. The antibiogram of one isolate showed sensitivity to ampicillin and all other antibiotics tested.[1]

Fourteen (64 percent) of the 22 cases studied were hospitalized for *S. arizona*-related illness. Two patients had multiple hospitalizations for their infections. One other case was hospitalized for unrelated congestive heart failure. The remaining seven cases were managed as outpatients. The median stay for *S. arizona* cases hospitalized once was 17 days.

Cases with *S. arizona* infection were significantly more likely to have an underlying chronic illness than *Salmonella* controls (OR = 6.7, CI = 1.1, 42.1). The most common chronic conditions among cases included AIDS (five cases), diabetes (four cases), and connective tissue disease (three cases); one case had both diabetes and connective tissue disease, and one case each had heart disease, cancer, and an illness of unknown etiology. Five of six *S. arizona*-infected cases (83 percent) without underlying disease had the bacteria isolated from feces only compared to one of 16 (6 percent) infected chronically ill patients with such an isolation pattern (95 percent CI of proportion difference = 0.33, 1.0). In unmatched analysis, controlling for chronic illness, cases were more likely than *Salmonella* controls to have an extragastrointestinal site of infection (OR = 26.3, CI = 1.4, 500). Three of six *S. arizona* patients without underlying

TABLE 1—Summary of Four Latino *S. arizona* Cases with Estimated Incubation Periods Greater Than Three Months, Los Angeles County, 1986–87

Case No	Date of Exposure to Rattlesnake Capsules	Date of First Positive Culture	Site of Positive Cultures	Chronic Illness	Comment
7	March 1987	8/13/87	Lung, urine	AIDS	Patient diagnosed with AIDS January 1987. Because the patient had chronic fever, the date of onset with regard to *S. arizona* is unclear.
17	1982	3/26/87	Perirectal Abscess	Adult onset diabetes mellitus and rheumatoid arthritis	Perirectal abscess first developed in June 1986 and recurred. Patient gave no history of symptoms which could be related to *S. arizona* infection prior to this time.
20	Early 1986	12/12/86	Perirectal Abscess	None	The date of onset is unclear. Patient took rattlesnake capsules for four months.
21	1985	9/8/86	Feces	None	Patient was hospitalized for 12 days during August 1986 for spinal surgery and received prophylactic antibiotics. Patient developed diarrhea during the hospitalization.

TABLE 2—Sites of Positive Cultures for *Salmonella arizona* by Disease Status in 22 Latino Patients, Los Angles County, 1986–87

Site of Isolate	No. Cases (%) Nomal Host (n = 6)	AIDS (n = 5)*	Other underlying conditions (n = 11)	Total (n = 22)
Feces	5 (83)	2 (40)	1 (9)	8 (36)
Blood		3 (60)	3 (27)	6 (27)
Lung or pleural fluid		2 (40)	2 (18)	4 (18)
Urine		1 (20)	2 (18)	3 (14)
Perirectal abscess	1 (17)		1 (9)	2 (9)
Wound			2 (18)	2 (9)
Lymph node		1 (20)		1 (5)

*One patient had positive blood and feces cultures, a second patient has positve lymph node, pleural fluid and feces cultures, and a third patient had positive lung and urine cultures.

illness (50 percent) were hospitalized with their infection.

Death occurred in six cases and one control. Three of the *S. arizona*-infected patients who died had AIDS. *S. arizona* infection was not listed as the direct cause of death in any of the cases. However, three patients had positive blood cultures for *S. arizona* within two months of their death and the infection appears to have exacerbated their condition.

Eighteen of 49 (37 percent) *S. arizona* isolates from the Latino cases were either serotype 3a $53:z_4,z_{23},z_{32}:-$(n = 9) or 3a $21:z_{29}:-$(n = 9) compared to none of 10 isolates serotyped from non-Latino patients in 1986 and 1987 (95 percent CI of proportion difference = 0.18, 0.57, Table 3). Both of these serotypes were cultured directly from rattlesnake capsules as were two other serotypes, 3a 61:1, v:1,5,7 and 3b 50:4:z, which were also isolated from both Latino and non-Latino patients. The 10 isolates of the latter two serotypes were all from feces, whereas 17 (94 percent) of the isolates of the former two serotypes were from extragastrointestinal sites.

Discussion

This investigation of Los Angeles County residents shows that human *S. arizona* cases are relatively common and that over three-fourths of recent *S. arizona* infections in Latinos are linked to rattlesnake capsules. Rattlesnake-capsule ingestion by persons with invasive *S. arizona* infection has been previously documented in case reports.[1–4] All of these cases were Latino patients who resided in the Southwest and had underlying disease.

Capsules cultured from *S. arizona* cases or households have been positive for *S. arizona* and other enteric bacteria.[1,2] The capsules are made of dried ground rattlesnake flesh and are known by several names including polvo de vibora, carne de vibora, or vibora de cascabel. The ingestion of snake preparations is an old popular folk remedy of Mexico and Central America and is used for a variety of medical conditions.[1,5] In this investigation, the majority of patients acquired the capsules from the border city of Tijuana, Mexico and from Los Angeles, California pharmacies or "farmacias" and "boticas" which serve the Latino community.

Reptiles constitute the major reservoir of *S. arizona*.[6–9] Other animal sources from which this bacterial species has been isolated are various fowl and rodents as well as opossums, goats, swine, dogs, and minks.[10,11] Foodborne outbreaks of *S. arizona* gastroenteritis have been reported associated with ice cream, cream pie, eggs, and egg powder.[12,13] Extragastrointestinal disease has been reported in patients with liver abscesses, septic arthritis, brain abscesses, pleural fluid infection, osteomyelitis, and cholecystitis.[12,14–16]

This surveillance study and case reports[1–4] indicate that underlying medical conditions are present in the majority of *S. arizona* cases who ingest rattlesnake capsules. Our findings suggest that *S. arizona* infections are at least as likely as other *Salmonella* serotypes to result in septicemia and focal

TABLE 3—*S. Arizona* Subgroup Serotypes by Ethnicity and Source of Isolation Los Angeles County, California 1986–87

Serotype	Number of Isolates Latino Extragastrointestinal	Intestinal	Non-Latino Extragastrointestinal	Intestinal	Rattlesnake Capsules
$G:z_4,z_{23}:-$	1				
$18:z_4,z_{23}:-$	1				
$18:z_4,z_{32}:-$	1		1		
$21:z_{29}:-$	8	1			1
$40:z_4,z_{23}:-$	3				
$40:z_4,z_{24}:-$		1			
$40:g,z_{51}:$			2		
$43:z_{24},z_{23}:-$		1		2	
$48:g,z_{51}:-$				1	
$48:z_{36}:-$					1
$48:i:z$	2	1			
$50:z_4,z_{24}:-$	1				
$50:k:z:z_{53}$				1	
$50:r:z$		2		2	1
$53:z_4,z_{23}:-$	1				
$53:z_4,z_{23},z_{32}:-$	9				1
$53:z_{29}:-$					1
$56:z_4,z_{23}:-$	3	1			
$60:r:e,n,x,z_{15}$		3			
$61:1:z$		1			
$61:1,v:1,5,7$		5		1	1
$65:(k):z$	1				
O rough:$z_4,z_{23}:-$	2				
O rough:$z_4,z_{23},z_{32}:-$					1
TOTAL	33	16	3	7	7

infections in patients with significant underlying disease.[17–19] The predisposing disease process alone or in association with treatments such as corticosteroids, chemotherapy, radiotherapy and antibiotics can interfere with host-resistance mechanisms. Edwards, *et al*,[12] have suggested, on the basis of isolation site proportions, that *S. arizona* may be more invasive than other *Salmonella* species. Our analysis using *Salmonella* controls and controlling for chronic disease is consistent with this possibility as is the 50 percent hospitalization rate in cases without underlying illness. Further, our data on *S. arizona* serotype distribution by isolation site and that of Weiss, *et al*,[20] suggest that invasiveness may be serotype-specific. In both studies, serotype 61:1,v:1,5,7 stands out as relatively noninvasive.

Most of the patients in this study ingested fewer than 10 rattlesnake capsules. Riley, *et al*,[1] found 10^5 *S. arizona* organisms per gram in one sample. Capsules weigh an average of 700 mg and, thus, the infectious dose of *S. arizona* may be as low as 10^5 or 10^6 organisms. For the many patients with reduced resistance who take rattlesnake capsules, the infectious dose may be even lower.

Although the incubation period for *Salmonella* gastroenteritis is usually 24–36 hours, four study persons apparently were indolently infected for months to years before developing signs of illness or before positive cultures were obtained. Prolonged intestinal carriage of *Salmonella* bacteria is a well recognized phenomenon; chronic illness or other factors such as surgery or disruption of intestinal flora due to antibiotics may have triggered active infection in these patients.[21] Two persons, one with underlying illness and one a normal host, had perirectal abscesses which could have been present for some time but were undetected. While recall and interviewer bias is possible, we believe the histories of rattlesnake capsule ingestion are accurate because interviewers described or denied such exposure without hesitation.

Five cases who were exposed to rattlesnake capsules had an underlying condition of AIDS. Investigators at Los Angeles County/University of Southern California Medical Center have recently reported on four of these patients.[1] Human immunodeficiency virus (HIV) infection has been shown to predispose patients to serious *Salmonella* infections.[22,23] Salmonellosis in persons with AIDS is characterized by widespread infection, bacteremia, and relapse even with standard antibiotic therapy. Through September 1988, over 900 cases of AIDS in Latinos have been reported in Los Angeles County, and the proportion of Latino AIDS cases has increased 133 percent from 9 percent to 19 percent between 1982 and 1988.* Thus, the dangers of rattlesnake-capsule exposure should be emphasized to Latinos due to their increasing pool of AIDS victims and the propensity of some to use folk remedies.

Physicians and health care workers should be aware of the potential for *S. arizona* infection from ingesting rattlesnake capsules and the prevalence of this practice. The State of California Department of Health Services Food and Drug Branch has taken action to prevent further distribution of the products and to remove the capsules from pharmacies and markets known to distribute them. Distributors were also requested to recall the products. Perhaps because of these actions and publicity in both Spanish and English, only 12 *S. arizona* cases linked to rattlesnake capsule ingestion were reported in Los Angeles County in 1988. Continued occurrence of cases is not surprising given the availability of rattlesnake pills in local Latino neighborhoods and along the US-Mexico border. Health agencies and community groups should pursue education of the Latino population about the risks and absence of medical benefit from rattlesnake capsules.

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Use of Medical Care for Chest Pain: Differences between Blacks and Whites

David S. Strogatz, PhD

Abstract: Data from a 1980, community-based survey of adult residents of Edgecombe County, North Carolina were analyzed to examine differences between Blacks and Whites in the reported use of medical care after experiencing chest pain. Of all adults (N=302) with chest pain in the year prior to interview, 49 percent of Blacks and 27 percent of Whites did not see a physician following the chest pain (difference = 22%, 95% CI = 12, 33). A multivariable analysis found that although the association between race and utilization was reduced at poverty levels of income, it was not explained by differences in demographic characteristics, health status or other dimensions of access to care. (*Am J Public Health* 1990;80:290–294.)

Introduction

Coronary heart disease (CHD) mortality rates for Black Americans exceed those of Whites at ages 25–64 and are similar to the rates of Whites for all ages combined.[1–3] In contrast, population rates of CHD occurrence based on cases from hospital discharge surveys are lower for Blacks than for Whites.[4,5] The inconsistency between the mortality and hospitalization data may be due to bias in detection and measurement, but it may also reflect an excess of CHD deaths among Blacks before the hospital is reached. A hospital surveillance study of CHD in Newark determined that Blacks were more likely to be dead on arrival at the hospital.[6] A study in urban and rural areas of South Carolina also reported that Blacks had higher rates of out-of-hospital mortality from acute myocardial infarction.[7] In community-based studies of sudden death, Blacks were found to be at greater risk than Whites in Charleston,[8] New Orleans,[9] and Nashville,[10] although not in Baltimore[11] where there was essentially no Black-White difference in the occurrence of sudden death.

Among those who do not die immediately, Blacks and Whites may differ by how soon medical care is received following the onset of symptoms. The National Ambulatory Medical Care Survey found that Blacks had fewer office visits for CHD and other diseases of the circulatory system.[12] Compared to Whites, Blacks are far less likely to be seen by a cardiologist, to have coronary angiography performed, and to undergo coronary artery bypass surgery.[12–15] These differences persist even among hospitalized patients, controlling for the clinical severity of disease.[14–17]

Less is known about Black-White differences in patients' knowledge of symptoms and decisions to seek care, an increasingly salient issue with the emergence of thrombolytic agents (e.g., streptokinase, tPA) for treatment in the early phase of myocardial infarction. A community survey on CHD conducted among Black residents of a southern US city found low levels of knowledge about CHD, its signs and symptoms.[18] In a comparison of Blacks and Whites living in Minneapolis-St. Paul, Blacks were less informed about the symptoms of myocardial infarction.[19] A study of attitudes about a variety of hypothetical symptoms found that Blacks and Whites had similar perceptions about how urgently to seek care for stable or unstable angina.[20] However, interviews with a consecutive series of Black patients hospitalized with acute myocardial infarction revealed that the average time between onset of symptoms and arrival at the hospital exceeded the times reported in similar investigations of White patients.[21] These preliminary findings indicate that gaps in knowledge and delays in receiving care may contribute to Black-White differences in pre-hospital mortality from CHD.

The purpose of this paper is to compare reported use of medical care by Blacks and Whites who experienced chest pain and to examine the degree to which demographic characteristics, access to medical care, and health status modified or accounted for Black-White differences in care-seeking behavior.

Methods

The data for these analyses were collected as part of the 1980 baseline survey of the Edgecombe County (North Carolina) High Blood Pressure Control Program. One thousand dwellings were randomly chosen, and interviews were sought with all members of the selected households who were 18 years of age or older. Data were collected from 2,029 individuals (1,085 White, 944 Black), representing an overall response of 91 percent. Additional information on the study population and procedures has been published.[22]

History of chest pain and associated use of medical care were measured by answers to a pair of questions from a larger set of similar items used to derive the symptom-response ratio, a need-based measure of utilization of health services.[23] During the interview, each individual was asked if he or she had experienced "repeated pains in or near the heart" at any time during the last 12 months and, if so, whether a doctor had been seen for it, in the past year or some time earlier. From this question, the 319 respondents (147 White, 172 Black) with chest pain in the past year were divided into three groups: 1) those who had seen a doctor for chest pain within the past year; 2) those who had seen a doctor for chest pain at some time, but not in the past year for this most recent episode; and 3) those who had never seen a doctor for their chest pain.

The baseline interview also included a series of questions concerning various barriers to medical care. As described in detail elsewhere,[24] these questions were used to measure four dimensions of access to care identified by Penchansky and Thomas[25]: *accessibility* (convenience of location of health services); *accommodation* (ease in getting an appointment); *affordability* (ability to pay for care); and *acceptability* (how well respondents get along with providers). Each dimension

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of access to care was represented by a structural score, a psychological score, or both. The structural access variables referred to rank ordering on tangible entities (e.g., distance to be traveled, time spent waiting, money or insurance at one's disposal), with a high score denoting good access. Psychological access referred to reported dissatisfaction on each dimension; scores on measures of psychological access represent the percent of individuals expressing dissatisfaction. The questions on access to care were answered by all persons who reported having a regular source of medical care. Fifteen of the 319 individuals with chest pain (3 White, 12 Black) did not have a regular source of care and therefore were not included in analyses involving these variables. Two other individuals had missing data on other variables of interest, leaving a total of 302 in the final analysis.

Other potential predictors of utilization which were measured at baseline and included in the analysis were age, sex, self-perceived health as measured by the Current Health Scale (CHS),[26] and reported history of myocardial infarction ("Have you ever had a heart attack?") and current antihypertensive treatment ("Are you now taking medicine for high blood pressure?")

The effect of race on the use of medical care for chest pain was initially characterized by an observed (unadjusted) difference between proportions with a corresponding confidence interval.[27] This measure conveys the potential impact of strategies to eliminate the association between race and utilization. Further analyses were conducted using linear multiple regression to estimate the Black-White difference adjusted for demographic characteristics, measures of health status, and dimensions of access to care. To highlight the contrast in use of care, these analyses were restricted to the two groups most different in utilization: individuals seen *during the past year* when chest pain occurred (group 1 defined above) vs individuals who were *not seen at all* for their chest pain (group 3). A backward stepwise procedure (with $p = .05$ to remain) was run to evaluate product terms combining race with each covariate.[27] Main effect terms for race and/for all covariates were automatically included in the model. With the linear regression analysis, the coefficient for the variable representing race was equal to the adjusted difference between the percent seeking care for Whites and the percent seeking care for Blacks. Race-specific regression models were also computed to assess the predictive importance of the covariates.

Results

The use of medical care by adults after they experienced chest pain is summarized in Table 1. Almost half of all Blacks with repeated pain near the heart never saw or discussed this with a doctor, compared to 27 percent of Whites (difference between percentages = 22%, 95% CI: 12, 33).

TABLE 1—Association between Race and Use of Medical Care by Adults Experiencing Chest Pain in the Past Year: Edgecombe County, NC, 1980

Use of Medical Care for Chest Pain	Black		White	
	N	(%)	N	(%)
MD seen in past year	35	(22)	67	(47)
MD seen, not in past year	45	(28)	38	(26)
MD never seen	78	(49)	39	(27)
Total	158	(100)	144	(100)

Subsequent analyses examined the distribution and influence of the covariates in addition to the Black-White difference in utilization. These analyses were restricted to individuals who either had seen a physician for chest pain in the past year (35 Blacks, 67 Whites—top row of Table 1) or had never seen a physician for chest pain (88 Blacks, 41 Whites). The characteristics of these adults are shown in Table 2. In the overall comparison of Blacks and Whites with chest pain, Blacks were younger, appeared to be healthier (by self-perception and history of myocardial infarction), had worse structural access to their source of medical care, and (except for cost) had somewhat more dissatisfaction with their source of care.

In the unadjusted data of both Blacks and Whites, utilization in response to chest pain was associated with older age, worse perceived health, and a history of myocardial infarction or current antihypertensive therapy. In contrast, crude associations between dimensions of access to care and utilization of care appeared to differ by race. Among Whites with chest pain, those adults who sought care had better structural access and less dissatisfaction with care. Among Blacks, care seeking was not associated with psychological access and had an inverse relationship with two dimensions (affordability, accommodation) of structural access.

The results of the multivariable analyses are given in Table 3. For Black adults, physician visits for chest pain were independently associated with history of myocardial infarction, current antihypertensive treatment, and having experienced delays in the past when trying to make an appointment (low structural accommodation). These variables were related to utilization in the same direction for Whites, but were less pronounced. The only predictor of significance for Whites was worse self-perceived health. Older age and relative satisfaction with the appointment process (psychological accommodation) were also associated with physician visits by Whites.

The predictors from the race-specific models were similarly identified in the model combining Blacks and Whites. The only statistically significant interaction term was the product of race and structural affordability (an index which combines information on income and insurance[24]). The race-specific models show that structural affordability was a weak positive predictor of utilization for Whites and negative for Blacks. The impact of this interaction on the Black-White difference in utilization can be illustrated for selected values of structural affordability (where the difference between percent utilization by Blacks and percent utilization by Whites equals $0.128 - 0.052 \cdot$ structural affordability). Among poor individuals (e.g., annual family income $\leq$ \$5,000, no health insurance: structural affordability = 4), Whites were more likely than Blacks to have seen a physician but the difference between the percentages was relatively small (8 percent). This difference between percentages of Whites and Blacks who saw a physician was 18 percent at the median value (6) for structural affordability in this sample. At a high level of resources for this community (e.g., annual family income = \$20,000, insurance pays for a portion of office visit and medication costs: structural affordability = 10) the difference was 39 percent.

Discussion

These data showed the expected Black-White difference in the use of medical care for chest pain. This difference increased with greater structural affordability (i.e., greater

TABLE 2—Characteristics of Adults Experiencing Chest Pain in the Past Year*: Edgecombe County, NC 1980

Characteristics**	Black Total (N = 113)	Black Care (N = 35)	Black No Care (N = 78)	White Total (N = 106)	White Care (N = 67)	White No Care (N = 39)
Sex (% male)	38	37	38	43	42	44
Mean Age (years)	42 (1.8)	50 (3.2)	38 (2.0)	51 (1.5)	55 (1.7)	43 (2.4)
Health						
Mean CHS	26.2 (1.0)	22.5 (1.6)	27.9 (1.2)	24.7 (1.0)	22.2 (1.3)	29.1 (1.6)
History of MI (% MI)	10	29	1	20	28	5
Treatment for HBP (% RX)	21	40	13	24	31	10
Mean Structural Access						
Affordability	6.1 (0.2)	5.7 (0.3)	6.2 (0.3)	7.7 (0.3)	8.0 (0.3)	7.3 (0.4)
Accommodation	7.3 (0.2)	6.9 (0.3)	7.5 (0.2)	7.5 (0.2)	7.7 (0.2)	7.1 (0.3)
Accessibility	10.8 (0.2)	10.9 (0.3)	10.7 (0.2)	11.5 (0.1)	11.6 (0.2)	11.3 (0.3)
Psychological Access (% dissatisfied)						
Affordability	53	51	54	58	51	72
Accommodation	48	46	49	41	30	59
Acceptability	51	49	49	41	36	49

*Restricted to (1) adults seen by a physician for chest pain during the past year and (2) adults never seen for chest pain.
**Mean and standard error reported for continuous variables, percent reported for dichotomous variables.

TABLE 3—Linear Regression Coefficients for the Multivariable Analysis of Race and Use of Medical Care by Adults Experiencing Chest Pain in the Past Year*: Edgecombe County, NC, 1980

Variables	Black β	Black SE	White β	White SE	Overall β	Overall SE
Sex	0.006	0.085	0.044	0.088	0.034	0.060
Age	0.001	0.003	0.006	0.003	0.003	0.002
Health						
CHS	−0.004	0.004	−0.009	0.005	−0.006	0.003
History of MI	0.585	0.143	0.148	0.125	0.333	0.091
Treatment for HBP	0.253	0.106	0.145	0.107	0.195	0.074
Structural Access						
Affordability	−0.026	0.019	0.024	0.018	0.080	0.038
Accommodation	−0.048	0.024	−0.017	0.027	−0.037	0.017
Accessibility	0.013	0.023	0.013	0.032	0.015	0.018
Psychological Access						
Affordability	−0.022	0.081	−0.104	0.094	−0.067	0.060
Accommodation	−0.074	0.085	−0.179	0.110	−0.139	0.067
Acceptability	0.045	0.083	−0.028	0.092	0.021	0.059
Race	—	—	—	—	0.128	0.184
Race × Structural Accessibility	—	—	—	—	−0.052	0.025

*Restricted to (1) adults seen by a physician for chest pain during past year and (2) adults never seen for chest pain.

family income and health insurance). Interpretation of these results requires recognition that measures of socioeconomic status may not have the same meaning and implications for Blacks and Whites.[28] For example, Blacks generally have had less opportunity than Whites to translate educational achievement into better jobs, more stable employment, and higher income,[29] so comparisons of Blacks and Whites with the same years of education can be misleading. Similar problems arise when comparing Blacks and Whites at the same level of income without adjustment for differences in family size, purchasing power, and property, all of which would tend to accentuate racial inequality in income.[30] The Black-White gap in utilization at higher levels of structural affordability in part may reflect differences in economic resources which may not be reflected by family income. These problems do not explain the inverse relationship between structural affordability and utilization among symptomatic Black adults. It may be a chance association in these data. Alternatively, social support is directly related to economic resources in this population[31]; perhaps the more economically advantaged Blacks rely more on social support networks for advice on health. Untapped dissatisfaction with past medical visits may also be reflected in lower utilization by those individuals with somewhat more income or insurance. As others have reported, the elimination of racial differences in the use of medical care will require more than just modifying the ability to pay for services.[32,33]

Other possible explanations for the Black-White difference in utilization include differences in perceptions about the seriousness of the symptom and the potential efficacy of medical care. Compared to Whites, Blacks may be less likely to perceive chest pain as serious enough to warrant immediate medical attention. Berkanovic and Telesky[34] suggest that the pronounced association between severity of symptoms and utilization for Blacks may be a vestige of past circumstances in which only the most disabling conditions could serve as justification for interruption of one's duties in order to seek care. This habit may persist today even for individuals who have more resources or autonomy over their personal lives. Studies of Whites with chest pain suggest that both the severity of the symptom and an individual's flexibility to respond at the time of symptom onset will contribute to delay in receiving care.[35,36]

Blacks and Whites also may view the likely efficacy of visiting the doctor differently. Perceived efficacy, an important predictor of care-seeking by symptomatic individuals,[37,38] was different in this population when treatment for hypertension was assessed.[24] Hypertensive Black adults were found to have less confidence in the net benefit of anti-hypertensive medication, and this attitude was associated with a greater likelihood of being untreated for diagnosed hypertension.

There are several limitations in these data. First, the question used to signify chest pain may not be measuring ischemic pain, and the meaning may vary by race. In Blacks and Whites, self-reported chest pain was associated with history of myocardial infarction and with established predictors of CHD (age, hypertension, and cigarette smoking), suggesting some common features of the measurement of the symptom. The instrument for the symptom-response ratio also included separate questions on indigestion and abdominal pain to avoid misclassification of chest pain in respondents with gastrointestinal symptoms.

Another problem is that the perceptions about health or access to health care may have been a product, as well as a determinant, of use of health services. Any resultant overestimation of the association between covariates and utilization (either because of selective recall or the influence of utilization on perceptions) would probably lead to an underestimate of the residual association between race and utilization. Despite this possibility, substantial Black-White differences were found in the multivariable analysis.

These findings require confirmation, e.g., in individuals from a defined population who present to the emergency room with acute myocardial infarction. The sequence of perceptions and behaviors from initial symptom to arrival at the hospital should be characterized. To the extent possible, similar data should be collected from witnessed instances of fatal CHD. Such a study was conducted 20 years ago in a White population[39]; replication in a Black or biracial community would be informative. The knowledge of symptoms of CHD also should be measured in community-based surveys related to cardiovascular disease.

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Award in Honor of Dr. Fred L. Soper (1893–1976) to Recognize Publications in the Field of Inter-American Health

A special award created by friends of Fred L. Soper, Director of the Pan American Health Organization (the World Health Organization Regional Office for the Americas) from 1947 to 1958, has been established in recognition of his outstanding contributions to health in the Americas.

In addition to his service with PAHO/WHO, Dr. Soper played a major role in the fight against yellow fever and other infectious diseases in Brazil as part of his work with the Rockefeller Foundation in the 1920s and 1930s and in the control of typhus in North Africa and Italy during the Second World War. He was one of the major figures of the century in inter-American health.

The award will be presented annually to the author or authors of an original scientific contribution comprising new information on, or new insights into, the broad field of public health, with special relevance to Latin America or the Caribbean or both. This may consist of a report, an analysis of new data, experimental or observational, or a new approach to analyzing available data. Preference will be given to studies involving more than one discipline and to papers related to infectious disease, a life-long concern of Dr. Soper. Review papers are eligible if their conclusions deal directly with policy primarily affecting Latin America or the Caribbean.

Papers that have been published or are suitable for publication in the *Boletín* of the Pan American Health Organization will be given priority, but meritorious papers published in other journals readily available to Latin American health personnel will also be eligible.

The award fund will be administered by the Pan American Health and Education Foundation (PAHEF) which will receive voluntary contributions designated for the purpose and hold them in a separate fund. The winner(s) of the award each year will be nominated by an awards committee consisting of experts designated by PAHO and PAHEF, and will be approved by the PAHEF Board of Trustees.

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Beliefs about AIDS, Use of Alcohol and Drugs, and Unprotected Sex among Massachusetts Adolescents

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Abstract: In August 1988, 1,773 Massachusetts 16–19-year-olds were surveyed by telephone using anonymous random digit dialing; response rate 82 percent. Logistic regression tested whether alcohol and drug use, perceived susceptibility to human immunodeficiency virus (HIV), severity of HIV if infected, effectiveness of condoms in preventing infection, barriers to condom use, and behavioral cues such as exposure to media or personal communication about acquired immunodeficiency syndrome (AIDS) were independently related to condom use. Among sexually active respondents, (61 percent of those interviewed) 31 percent reported always using condoms. Respondents who believed condoms are effective in preventing HIV transmission and worried they can get AIDS were 3.1 and 1.8 times, respectively, more likely to use condoms all the time. Respondents who carried condoms and who had discussed AIDS with a physician were 2.7 and 1.7 times, respectively, more likely to use them. Those who believed condoms do not reduce sexual pleasure and would not be embarrassed if asked to use them were 3.1 and 2.4 times, respectively, more likely to use condoms. Teens who averaged five or more drinks daily or used marijuana in the previous month were 2.8 and 1.9 times, respectively, less likely to use condoms. Among respondents who drink and use drugs, 16 percent used condoms less often after drinking and 25 percent after drug use. Those counseling adolescents about HIV should assess and discuss beliefs outlined in the Health Belief Model, as well as their alcohol and drug use. (*Am J Public Health* 1990; 80:295–299.)

Introduction

One-fifth of reported AIDS (acquired immunodeficiency syndrome) cases in the United States have been in the 20–29 year age group.[1] Given the incubation period, many of those persons were infected during their high school years or shortly thereafter.

Knowledge of behavioral risks for AIDS does not prompt all at risk to change their behavior.[2] Most older teens in Massachusetts have learned the principle modes of human immunodeficiency virus (HIV) transmission, and adolescent condom use has increased since 1986.[3,4] However, two-thirds of sexually active 16–19 year olds surveyed there in the fall of 1988 reported sexual intercourse in the past year without the use of condoms.[4] Nationally, teenagers account for one-fourth of sexually transmitted disease cases.[5]

Educators need to move beyond imparting knowledge about modes of HIV transmission to change adolescent behavior. A specific set of beliefs described in the Health Belief Model has been found to predict behaviors to avoid other health risks,[6] and has been hypothesized to also predict whether someone will adopt behavior to avoid HIV transmission.[7] Persons who believe they are personally susceptible to HIV exposure, the consequences of exposure are severe, protective measures such as condom use are effective in preventing transmission, and perceive few barriers to the adoption of condom use may be more likely to adopt such behavior to avoid HIV transmission. According to this model, people make a rational cost-benefit analysis when trying to decide whether to adopt preventive behavior. Actual changes in behavior may then be stimulated by cues to action such as educational messages or learning that someone they know has AIDS.

However, among adolescents, sexual encounters may be unplanned and also often occur after drinking or drug use. Exploratory studies in San Francisco indicate gay men are more likely to have unprotected intercourse after alcohol and drug use.[8] Whether this applies to heterosexuals warrants exploration. Because these and other factors may mitigate against adolescents making a cost-benefit analysis about HIV and condom use before each sexual encounter, it is important to assess whether the beliefs described in the model do indeed predict which adolescents will use condoms and, if so, which beliefs are the strongest independent predictors of use. It is also important to assess whether alcohol and drug use are negatively related to condom use and, if so, why.

This Massachusetts statewide study of adolescents assessed:

- whether beliefs outlined in the Health Belief Model are significantly associated with condom use among sexually active adolescents;
- which beliefs are the strongest independent predictors of condom use after controlling for social and demographic characteristics of respondents;
- whether condom use is less likely after drinking and drug use; and
- whether those who use condoms less often after drinking and drug use have different characteristics and beliefs about HIV transmission.

Methods

In September 1988, a statewide anonymous random digit dial telephone survey of 16-to-19-year-olds was conducted in Massachusetts using methods developed by Waksberg.[9] Within each household, one adolescent was randomly selected using methods developed by Kish.[10] The survey of 1,773 adolescents (response rate 82 percent) was conducted in English and Spanish and explored respondent knowledge and beliefs about HIV transmission; alcohol, marijuana, and other illegal drug use; and behavior change in response to the AIDS epidemic. The study specifically asked respondents beliefs that are identified in the Health Belief Model (see Appendix for details).

In addition, respondents were asked their age, level of education, gender, whether their sexual activity was with males, females or both, and the number of sexual partners

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they had in the past year. Chi square analyses tested whether individual beliefs outlined in the Health Belief Model and alcohol and drug use were associated with condom use. Condom use was classified as: never used condoms, use condoms some of the time, always used condoms. Stepwise logistic regression analysis[11] tested which beliefs were the strongest independent predictors of condom use after considering demographic characteristics, alcohol and drug use, and numbers of sexual partners, and whether alcohol and drug use were independently related to condom use. Condom use was dichotomized as always used condoms versus sometimes or never used. Odds ratios and confidence intervals were calculated. The chi square test for symmetry[12] assessed whether individuals were less likely to use condoms after drinking and drug use. Finally, chi square analyses tested for differences in characteristics and beliefs of adolescents who used condoms less often after drinking or drug use.

Results

Bivariate Analysis

Condom use according to respondent characteristics—Among the 1,773 adolescents interviewed, 61 percent (n = 1,050) reported having had intercourse in the previous year. Two percent of males reported homosexual behavior. Among sexually active teens, 31 percent (n = 320) reported always using condoms, 32 percent (n = 345) sometimes, and 37 percent (n = 385) never used condoms. Males were more likely than females to report always using (34 percent vs 26 percent). Thirty-nine percent of 19-year-olds compared to 29 percent of 16-year-olds were more likely to always use condoms. Thirty-one percent of Whites, 28 percent of Blacks, and 39 percent of Hispanics, but only 7 percent of persons from other racial backgrounds, always used condoms. Condom use varied by previous numbers of sexual partners (Table 1). Respondents who reported only one partner in the last year were much more likely to never use condoms than those who had more partners. However, respondents who reported 10 or more partners were the least likely to always use condoms.

TABLE 1—Condom Use Among Sexually Active Massachusetts Teenagers According to Numbers of Sexual Partners

Number of Sexual Partners (past year)	% Never Use Condoms N = 380	% Sometimes Use Condoms N = 337	% Always Use Condoms N = 316	
1 (N = 549)	46	26	29	
2–4 (N = 370)	29	35	36	(p<.0001)
5–9 (N = 85)	22	52	26	
10+ (N = 28)	8	75	17	

Adolescents who averaged five or more drinks daily were somewhat less likely to always use condoms than abstainers (29 vs 35 percent) as were persons who smoked marijuana at least weekly (22 vs 32 percent of abstainers). Persons who used other psychoactive drugs in the past month were also less likely to always use condoms than those who did not (20 vs 32 percent).

Condom use according to respondents' health beliefs—On most dimensions of the Health Belief Model, bivariate associations with condom use were observed. Respondents were more likely to always use condoms if they felt susceptible to AIDS; believed condoms are effective; perceived few barriers to condom use; and were exposed to more cues to action (Tables 2–4).

Multivariate Analysis

Health beliefs—We used stepwise logistic regression analysis to assess which beliefs were the strongest independent predictors of whether sexually active adolescents always used condoms (Table 5). Variables under barriers to use were the strongest predictors of condom use. Perceived

TABLE 2—Proportion of Massachusetts Adolescents Who Always Use Condoms According to Their Perceived Susceptibility to HIV and AIDS (N = 1,050)

Perceived Susceptibility	N	% Always Use Condoms	Relative Risk*	(95% C.I.)*
How likely is a female to acquire the AIDS virus during intercourse with an infected male?				
Very likely	(749)	30	9.44	(1.37,65.09)
Somewhat likely	(270)	34		
Not likely/a little likely	(31)	3		
How likely is a male to acquire the AIDS virus during intercourse with an infected female?				
Very likely	(688)	30	1.70	(1.01,2.88)
Somewhat likely	(293)	35		
Not likely/a little likely	(68)	18		
How much do you worry you could get AIDS?				
A great deal	(161)	47	1.81	(1.46,2.26)
Somewhat	(357)	31		
A little/not at all	(532)	26		
Perceived Severity				
Likelihood an infected person will die of AIDS?				
Very	(868)	31	.53	(.32,.86)
Somewhat	(170)	27		
Not at all/a little	(12)	58		
Is there a cure for AIDS?				
No	(1,035)	31	4.62	(.69,30.78)
Yes	(14)	7		

*When a variable has more than two categories the extreme categories were compared.

TABLE 3—Proportion of Sexually Active Massachusetts Adolescents Who Always Use Condoms According to Their Perceived Effectiveness and Barriers to Use (N = 1,050)

Effectiveness	N	% Always Use Condoms	Relative Risk*	(95% C.I.)*
Perceived Effectiveness of Condoms in Reducing HIV Transmission				
Very/extremely	(693)	31	5.50	(1.43,21.24)
Somewhat	(321)	31		
Not at all/a little	(35)	6		
Barriers to use				
Carry condoms				
Yes	(267)	49	2.06	(1.73,2.45)
No	(783)	24		
Condoms reduce pleasure				
Not at all true/a little	(484)	40	3.19	(2.16,4.71)
Somewhat true	(367)	27		
Very true	(189)	13		
Condoms are difficult and embarrassing to obtain				
Not at all true	(384)	31	1.29	(.92,1.82)
Somewhat true/a little	(537)	31		
Very true	(128)	24		
Would be embarrassed if a partner asked condoms to be used				
Not at all true	(798)	33	2.06	(.98,4.31)
Somewhat true/a little	(208)	23		
Very true	(37)	15		

*When a variable had more than two categories the extremes were compared.

effectiveness of condoms, and how much respondents worried about acquiring AIDS also independently predicted use. Also, adolescents who spoke to physicians about AIDS were more likely to use condoms.

Of the demographic and behavioral variables entered into the same regression analysis, age predicted always using condoms, with younger persons more likely to use condoms. Race was also predictive of use, with Whites more likely than Blacks or other non-Whites to use condoms. Independent of beliefs, number of sexual partners, education, and gender were not related to condom use.

TABLE 4—Proportion of Sexually Active Massachusetts Adolescents Who Always Use Condoms According to Their Exposure to Cues to Action (N = 1,050)

Cues to Action	N	% Use Condoms	Relative Risk	(95% CI)
Know someone with AIDS				
Yes	(57)	16	.50	(.27,.92)
No	(993)	31		
Discussed AIDS with friends				
Yes	(920)	32	1.48	(1.05,2.08)
No	(130)	21		
Discussed AIDS with Physician				
Yes	(161)	46	1.66	(1.36,2.02)
No	(889)	28		
Media Exposure to AIDS information				
Yes	(701)	32	1.21	(.99,1.49)
No	(349)	27		
Teacher discussed about AIDS				
Yes	(778)	31	1.10	(.86,1.40)
No	(197)	28		

TABLE 5—Logistic Regression of Behaviors and Beliefs Associated with Consistent Condom Use among Sexually Active Massachusetts Adolescents

Health Belief Variable	Odds Ratio	(95% CI)
Susceptibility-Severity		
Worry can get AIDS (great deal vs not at all)	1.82	(1.13,2.92)
Effectiveness		
Condoms effective in preventing HIV transmission (extremely vs not)	3.12	(1.37,7.14)
Barriers to Use		
Condoms reduce pleasure (not true vs very true)	3.13	(1.94,5.04)
Carry condoms (yes vs no)	2.73	(1.92,3.89)
Embarrassed if asked to use condom (not vs very)	2.36	(1.16,4.79)
Cues to Action		
Discussed AIDS with Physician	1.67	(1.13,2.47)
Other Characteristics and Behaviors		
Average daily drinks (0 vs 5+)	2.77	(1.42,5.41)
Days smoked marijuana (0 vs 5+ past month)	1.86	(1.06,3.25)
Age (16 vs 19 years)	2.48	(1.56,3.94)
Race:		
White vs Black	2.13	(1.25,3.64)
White vs Other	2.91	(1.02,8.27)

Drinking and drug use—Adolescents who drank five or more drinks daily (5 percent of sexually active adolescents) and those who used marijuana in the past month (29 percent of sexually active adolescents) were 2.8 and 1.9 times, respectively, less likely to use condoms. To further assess whether drinking and drug use independently influenced condom use, we asked respondents if they had intercourse in the past six months after drinking and after drug use. Those who had intercourse were then asked how often they used condoms when having sex after drinking and after drug use, and how often they used condoms when they had not been drinking or taking drugs.

Sixteen percent of adolescents who said they had sex after drinking reported using condoms less often after drinking than when not drinking (73/465) (Table 6) and 25 percent who said they had sex after drug use reported using condoms less often after using drugs (40/160) (Table 7).

Compared to other sexually active adolescents who drink, those who use condoms less often after drinking were significantly more likely to believe HIV sexual transmission, if one partner is infected, is only a little likely from males to females (10 vs 2 percent), females to males (19 vs 2 percent), or between males (5 vs .1 percent). They were also more likely to believe it is very true that condoms reduce pleasure (27 vs 15 percent), to have sex after drug use (45 vs 28 percent), and to have had 10 or more sexual partners in the past year (11 vs 4 percent).

Compared to other sexually active adolescents, those

TABLE 6—Condom Use When Drinking and not Drinking among Sexually Active Massachusetts Teenage Drinkers

Condon Use When Drinking	Condom Use When not Drinking					
	Always		Sometimes		Never	
	N	(%)	N	(%)	N	(%)
Always	194	(42)	17	(4)	8	(2)
Sometimes	40	(9)	140	(30)	11	(2)
Never	4	(1)	29	(6)	22	(5)

Use condoms less often after drinking: 73/465, 16%

TABLE 7—Condom Use When Taking Drugs and not Taking Drugs among Massachusetts Teenagers Who Use Drugs

Condom Use When Taking Drugs	Condom Use When not Taking Drugs					
	Always		Sometimes		Never	
	N	(%)	N	(%)	N	(%)
Always	43	(27)	1	(1)	—	—
Sometimes	25	(15)	61	(37)	15	(9)
Never	—		15	(9)	—	—

Use condoms less often after taking drugs 40/160, 25%

who used condoms less often after other psychoactive drug use also more often believed it very true that condoms reduce pleasure (34 vs 16 percent). They tended to be younger (45 vs 21 percent age 16), more likely to have dropped out of school (25 vs 8 percent), and much more likely to believe HIV is transmitted in ways it is not, e.g., by sharing eating and drinking utensils, and shaking hands (both 9 vs 1 percent).

Discussion

Telephones are owned by over 95 percent of households in Massachusetts, but low-income racial/ethnic minority households are the least likely to own phones. Because low-income households are at the highest risk of AIDS, this survey may slightly overestimate respondents' knowledge about AIDS and underestimate prevalence of behaviors that pose risk of HIV transmission.

A study of adults in Montreal found beliefs in the Health Belief Model predicted safer sex practices among adults.[3] This study suggests that some beliefs about HIV transmission and AIDS identified in the Health Belief Model predict whether adolescents will use condoms. Concern about acquiring AIDS, beliefs that condoms are effective in preventing HIV transmission, and speaking to a physician about AIDS predicted use. Barriers to use such—i.e., beliefs that condoms reduce pleasure, are embarrassing to use, or that respondents do not routinely carry condoms—were also highly predictive of not using condoms.

Variables concerning perceived severity of HIV infection were not, however, significant independent predictors of use. Although bivariate analyses revealed these variables were related to condom use, the overwhelming majority of respondents believed there is no cure for AIDS, and that infected persons will die from AIDS. Consequently, those variables did not explain much variation in condom use.

Beliefs about AIDS accounted for approximately one-tenth of variation in use. While a sizable fraction, this means that numerous other issues relate to use. In this study, use was less among heavy drinkers and frequent drug users suggesting the limits of rational cost-benefit beliefs in determining condom use when people are intoxicated.

A study of adolescents in Scotland found that respondents' first sexual intercourse was more likely to have occurred without the use of a condom if they had been drinking.[14] In this Massachusetts study, even among the same individuals, use of condoms were less likely after drinking and drug use suggesting that concerns about condom use may seem less pressing when people are intoxicated.

Many other factors not explored in this survey and not directly related to health beliefs about condoms and AIDS or HIV may also influence whether adolescents use, or will use, condoms. These include the ability to discuss contraception with sexual partners; perceptions of which partner's responsibility it is to purchase and use contraceptives; beliefs about abortion; concerns about other sexually transmitted diseases; beliefs and desires for pregnancy; beliefs about the efficacy of condoms relative to other contraceptives; issues of ambivalence about being prepared for sexual intercourse; influence of peers and perceptions about whether condom use is increasing among peers; advice received about family planning; perceptions of medical care providers and the adequacy or sensitivity of their advice, and cultural and religious beliefs. Beliefs and preferences of partners on all the issues cited above could also influence condom use. Also, although 77 percent of adolescents reported that they would not be embarrassed if a partner asked that condoms be used, this was an abstract response to a question most had never been asked. If one sex partner asks another to use a condom, questions of trust and fidelity could easily surface for both partners.

Nonetheless, this study suggests that the Health Belief Model provides a useful set of beliefs about AIDS to explore among adolescents and to target in efforts to increase condom use. Not only should educators try to change beliefs through education about personal susceptibility to the virus,[10] they should try to heighten perceived effectiveness of condoms by teaching adolescents which types of condoms offer the most protection, and how to avoid mistakes in using condoms that could reduce their effectiveness. Issues of embarrassment should be addressed possibly by role playing of discussion with sexual partners about AIDS and condoms, and efforts could be undertaken to make condoms more easily accessible to both male and female adolescents.

Finally, alcohol and drug use may reduce the likelihood that adolescents will use condoms. The array of educational, counseling, and legal approaches to reduce adolescent alcohol and drug consumption may thus have a particularly important influence on HIV transmission.

APPENDIX

Questions Used in Massachusetts Survey of Beliefs, Knowledge about AIDS/HIV and Behavior

1) *Susceptibility*
    - How much do you worry that you could get AIDS in your lifetime? (A great deal, Some, A little, Not at all.)
    - How likely is it that if a man and woman have sex, and the man is infected with HIV, that the woman will become infected? (Very, Somewhat, A little, Not at all.)
    - How likely is it, if the woman is infected, the man will become infected? (Very, Somewhat, A little, Not at all.)
    - If two men have sex and one is infected how likely is it the other will become infected? (Very, Somewhat, A little, No at all.)
2) *Severity*
    - Is there a cure for AIDS? (Yes, No)
    - How likely is it that someone with AIDS will die from the disease? (Very, Somewhat, A little, Not at all.)
3) *Effectiveness*
    - How effective are condoms in preventing transmission of AIDS virus? (Extremely, Very, A little, No at all effective.)
4) *Barriers to Action*
    - How true do you think it is that:
        - — condoms are difficult and embarrassing to obtain (Very, Somewhat, A little, Not at all)
        - — condoms reduce pleasure (Very, Somewhat, A little, Not at all)
    - If your partner wanted a condom to be used during intercourse, how likely is it that:
        - — you would be embarrassed (Very, Somewhat, A little, Not at all)
        - — you would be upset with your partner (Very, Somewhat, A little, Not at all)

- Do you usually carry condoms with you? (Yes, No)

5) *Cues to Action*

- Did you read or hear anything about AIDS, in the past week, in the newspaper, magazine, radio, or T.V.? (Yes, No)
- Have you ever discussed AIDS with a parent or other adult member of your family? (Yes, No)
- Have any of your teachers in school every talked about AIDS? (Yes, No)
- Have you ever discussed AIDS with a physician (Yes, No)
- Do you know anyone with AIDS? (Yes, No)

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AAFP and CDC Recommend Second MMR Vaccination at Preschool Age

The American Academy of Family Physicians (AAFP) recommends routine second measles, mumps and rubella (MMR) immunizations for children four to six years of age, or just prior to entering school. The national association representing more than 68,000 family doctors, residents, and medical students also said that children in areas where there is a high risk of contracting measles should receive their first MMR at 12 months, instead of 15 months of age. The revised guidelines calling for two childhood measles vaccinations concur with the Immunizations Practice Advisory Committee's report: *Measles Prevention*, issued recently by the federal Centers for Disease Control (CDC). The AAFP revised its schedule for MMR vaccinations because of the increasing number of infants, children and adolescents who have contracted measles. About five out of every 100 children previously immunized failed to develop protection against measles.

The new guidelines define a high-risk area as a county or portion of a county with:

- more than five measles cases among preschool-aged children during each of the last five years;
- a recent outbreak among unvaccinated preschool-aged children; or
- a large urban population.

CDC also recommends a second MMR shot or a measles immunity test for everyone who is going to college or other educational institutions beyond college, as well as medical personnel beginning employment who will have direct patient care contact.

Persons who are *not recommended* to receive a measles vaccination generally include pregnant women, persons with active tuberculosis, or those who have allergies to eggs.

The Growing Proportion of Female Physicians: Implications for US Physician Supply

Phillip R. Kletke, PhD, William D. Marder, AM, and Anne B. Silberger, AM

Abstract: This study analyzes how the growing proportion of women in the United States physician population will affect the amount and type of physician services available to the US population. Female physicians work fewer hours per week, are slightly less likely to be in patient care, and tend to enter different specialties than male physicians. Female physicians also have higher retirement rates than male physicians, but due to their lower mortality rates, have work lives nearly as long as male physicians. We examined how the changing composition of the physician population will affect the availability of physician services by comparing historical and projected trends for the number of active post-residency physicians with comparable trends for a full-time-equivalent measure of physician supply. The full-time-equivalent measure takes into account the different labor supply behavior of key subpopulations (e.g., women and graduates of US versus foreign medical schools). The results suggest that the changing composition of the physician population will reduce the growth of effective physician supply between 1986 and 2010 but only by four percentage points. (*Am J Public Health* 1990; 80:300–304.)

Introduction

The proportion of United States physicians who are women has grown rapidly since 1970 and, according to recent projections,[1] it will continue to rise until the year 2010. Since the practice characteristics of female physicians differ from those of male physicians, the growing representation of women in the medical profession is likely to have significant effects on the future supply of physicians.

In general, women physicians have been disproportionately represented in pediatrics and psychiatry and have been underrepresented in the surgical specialties.[2,3] Analyses of the most recent cohorts of new physicians, however, show that the specialty distributions of male and female physicians appear to be converging.[4] This convergence is much more pronounced among graduates of US medical schools (USMGs) than graduates of foreign medical schools (FMGs).[5] Nevertheless, there are still significant differences in the specialty choices of male and female US medical students.[6]

Numerous studies have documented that female physicians work fewer hours per week than male physicians,[7-11] and several analyses have explored the extent to which this difference is attributable to marital status and household responsibilities.[12,13] Longitudinal data show that the differences in the number of hours male and female physicians work may be decreasing, although significant differences still remain.[14] Because women physicians work fewer hours per week than male physicians and are comprising an increasing proportion of the physician population, some analysts have concluded that current projections of the physician population overstate effective future physician supply.[15,16] Partly in response to this issue, a recent report of the US Council on Graduate Medical Education[17] indicated that the practice patterns of male and female physicians deserve greater research. This paper measures current differences between male and female physicians and traces their consequences for future physician supply.

Methods

We used the Physician Masterfile[2] of the American Medical Association (AMA) for both current and historical data on all US physicians (both AMA members and nonmembers). We used the Demographic Model of the Physician Population[1] to project the size and composition of the future physician population.

Our projection of the physician population assumed that:

- the proportion of women among graduates of US medical schools will increase 1 percent per year from 32 percent in 1986 until it reaches 40 percent and will remain constant thereafter;
- male and female physicians will choose specialties following patterns identical to the most recent cohorts of medical school graduates; and
- retirement and mortality rates for male and female physicians will remain constant at their current levels.

We analyzed data from the AMA Socioeconomic Monitoring System[11] to examine practice differences of male and female physicians. These data come from a series of periodic telephone surveys of random samples of non-federal, post-resident, patient care physicians since 1982. We used data from the 1987 core survey, which has data on approximately 4,000 physicians.

Results

Demographic Characteristics

Table 1 shows that the proportion of women physicians increased from 7.1 percent to 15.3 percent between 1970 and 1986, and is projected to reach almost 30 percent by the year 2010. Consequently, women physicians are concentrated in the younger age categories. In 1970 and 1986, female physicians were disproportionately represented among foreign medical graduates not born in the US, but the relationship between sex and country of graduation reverses during the projection period due to the growing number of females graduating from US medical schools. Female physicians are concentrated in pediatrics and psychiatry and are underrepresented in general surgery and the surgical subspecial-

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TABLE 1—Proportion of Active Physicians who are Female According to Age, Specialty, and Country of Graduation for Years 1970, 1986, 2000, and 2010

	Historical		Projected	
	1970	1986	2000	2010
Total	7.1%	15.3%	24.1%	29.4%
Country of Graduation				
USMG	5.4	14.0	24.7	30.7
FMG	14.6	20.0	21.7	23.0
USFMG	4.2	10.7	13.0	14.0
Foreign-born FMG	16.0	21.7	24.0	25.9
Age (years)				
< 36	9.5	24.5	36.5	37.3
36–45	6.4	15.6	30.3	37.1
46–55	6.3	9.5	20.5	30.9
56–65	5.0	7.1	12.8	21.4
> 65	5.4	5.9	8.8	13.8
Specialty				
General/Family Practice	4.4	12.9	22.9	28.8
General Internal Medicine	5.8	17.4	26.5	31.5
Medical Subspecialties	5.0	10.2	18.4	22.8
General Surgery	1.1	5.8	10.6	14.0
Surgical Subspecialties	1.3	3.9	6.8	8.9
Pediatrics	21.4	35.3	47.5	54.1
Obstetrics/Gynecology	7.2	19.6	35.9	45.6
Radiology	5.3	12.1	19.7	24.4
Psychiatry	13.2	22.4	34.0	41.8
Anesthesiology	14.3	17.4	23.2	26.6
Pathology	12.9	21.8	29.2	34.7
Emergency Medicine	—*	12.9	20.5	24.3
Other Specialties	9.6	18.1	26.4	31.7

SOURCE: AMA Physician Masterfile, 1970–86, Department of Data Release Services, Division of Survey and Data Resources; AMA Demographic Model of the Physician Population, AMA Center for Health Policy Research, 1988.

ties. These patterns appear in both the 1970 and 1986 data as well as the projection estimates for 2000 and 2010.

Differences in the specialty distributions of male and female physicians persist but are gradually weakening, however. The index of dissimilarity for the specialty distributions of male and female post-resident physicians (i.e., the percent of female physicians who would have to change specialties so that their specialty distribution would match that of male physicians) dropped from 31.2 to 23.7 between 1970 and 1986, indicating that the concentration of female physicians in certain specialties has decreased. The index calculated only for physicians less than 35 years of age dropped from 28.5 to 18.4 during the same period.

The mortality, retirement, and accession (resumption of activity) rates for the projection model were estimated from an analysis of Masterfile data for years 1978 to 1986 from a set of logistic regression equations, based on the physician's age, sex, specialty, and country of graduation. (Analyses available on request to author.) Female physicians had higher retirement rates and lower mortality rates than male physicians. By assuming that the schedules of transition rates remain constant, we estimated how much longer physicians of a given age, sex, specialty, and country of graduation are, on average, likely to live, and how much longer they are likely to be professionally active.

Table 2 presents the expected years of remaining life and the expected years of remaining work life for active male and female USMG physicians at ages 35, 50, and 65 in 13 different specialty groups. Physicians on average, live and work longer than the rest of the US population. Female physicians can expect to live approximately four to five years longer than male physicians, but their higher retirement rates mean that active female physicians at age 35 have approximately the same expected work life as male physicians at 35.

TABLE 2—Expected Life and Work Life of Active USMG Physicians at Ages 35, 40, and 65 by Specialty and Sex

	Age (years)					
	35		50		65	
	Male	Female	Male	Female	Male	Female
Expected Life						
Specialty General/Family Practice	42.4	47.2	28.4	33.1	16.3	20.[illegible]
General Internal Medicine	43.8	48.4	29.7	34.2	17.3	21.3
Medical Subspecialties	45.0	49.6	30.8	35.3	18.2	22.2
General Surgery	43.7	48.4	29.6	34.1	17.2	21.2
Surgical Subspecialties	44.5	49.2	30.4	34.9	17.9	21.9
Pediatrics	44.9	49.6	30.8	35.3	18.2	22.2
Obstetrics/Gynecology	43.1	47.7	29.0	33.6	16.8	20.7
Radiology	44.4	49.2	30.2	34.9	17.8	21.9
Psychiatry	43.4	48.2	29.3	34.0	17.1	21.2
Anesthesiology	43.9	48.6	29.8	34.4	17.4	21.4
Pathology	43.6	48.6	29.6	34.4	17.3	21.5
Emergency Medicine	42.7	47.9	28.8	33.8	16.8	21.0
Other Specialties	43.7	48.6	29.6	34.4	17.4	21.5
Expected Work Life						
Specialty						
General/Family Practice	36.5	36.8	22.7	23.5	11.2	12.3
General Internal Medicine	37.6	37.8	23.6	24.3	11.8	13.0
Medical Subspecialties	39.1	39.7	25.1	26.1	13.0	14.3
General Surgery	37.1	37.0	23.1	23.7	11.5	12.5
Surgical Subspecialties	37.7	37.5	23.6	24.1	11.8	12.8
Pediatrics	37.8	37.6	23.7	24.2	11.9	12.9
Obstetrics/Gynecology	37.0	37.0	23.0	23.6	11.4	12.4
Radiology	36.5	36.0	22.6	22.8	11.0	11.9
Psychiatry	37.5	37.9	23.5	24.5	11.8	13.1
Anesthsiology	35.8	34.9	21.9	21.9	10.5	11.2
Pathology	36.3	35.9	22.4	22.8	10.9	11.9
Emergency Medicine	34.9	34.2	21.1	21.3	10.0	10.8
Other Specialties	35.8	35.0	22.0	22.1	10.6	11.4

SOURCE: Marder, *et al. Physician Supply and Utilization by Speciality: Trends and Projections.* Table 3.1, p 42.

Physician Time in Patient Care

Among the entire post-residency physician population, 89.7 percent of male physicians and 88.0 percent of female physicians are classified as being in direct patient care (i.e., spending the plurality of their professional working hours treating patients in either an office-based or hospital-based practice). The proportion decreases somewhat for older physicians, and both USFMGs and foreign-born FMGs are slightly more likely than USMGs to be in patient care. (Data available upon request to author.) The probability of being in patient care was highest in general/family practice, the surgical subspecialties, and obstetrics/gynecology, and lowest in the medical subspecialties, pathology, and in the residual "other specialties" category. There are few differences by gender.

Table 3 shows that post-residency patient care physicians worked an average of 58.3 hours per week in 1987. On average, female physicians, older physicians and USMGs worked fewer hours than their counterparts. Physicians in general internal medicine, the medical subspecialties, pediatrics, obstetrics/gynecology, and emergency medicine worked several more hours per week than average, while physicians in psychiatry, pathology, and the other specialties worked fewer hours than average.

TABLE 3—Average Hours Worked per Week by Male and Female Non-Federal Post-residency Patient Care Physicians According to Age, Specialty, and Country of Graduation, 1987

	Total	Male	Female
Total	58.3	59.0	52.3
Country of Graduation			
USMG	58.1	58.6	52.3
FMG	59.3	60.6	52.2
USFMG	59.1	59.3	55.0
Foreign-born FMG	59.3	60.9	52.1
Age (years)			
< 36	60.3	61.7	53.9
36–45	60.2	61.4	50.6
46–55	58.8	59.2	54.3
56–65	55.6	55.8	51.9
> 65	49.0	49.1	45.7
Specialty			
General/Family Practice	58.2	59.3	48.0
General Internal Medicine	60.6	62.0	50.5
Medicine Subspecialties	62.1	62.4	56.1
General Surgery	59.7	59.7	53.7
Surgical Subspecialties	57.8	57.9	52.5
Pediatrics	59.8	62.2	53.3
Obstetrics/Gynecology	60.9	61.7	56.2
Radiology	57.2	57.3	54.8
Psychiatry	50.4	51.1	46.5
Anesthesiology	58.8	60.3	49.7
Pathology	52.9	52.5	56.1
Emergency Medicine	60.7	60.2	64.3
Other Specialties	54.6	54.7	54.4

SOURCE: Marder, *et al.* *Physician Supply and Utilization by Specialty: Trends and Projections.* Table 2.6, p 18.

The findings reported for the proportion of physicians in patient care and their average hours of work per week were confirmed in a logistic regression analysis of whether a physician was in patient care and an ordinary least squares regression analysis of the log of hours worked per week by patient care physicians. (Analysis available on request to author.) Independent variables were: the physician's sex, country of graduation, specialty, physician's age and age-squared (since the relationships with the dependent variables did not appear to be linear with respect to age), and interaction terms for sex with age and age-squared.

In each analysis, the coefficients for sex and the age-sex interactions were jointly significant at the .001 level. Plots of the age-specific regression estimates for male and female physicians indicated that female physicians are slightly less likely to be in patient care than male physicians for all ages. The hours worked per week by patient care physicians peak at age 37 for male physicians and at age 43 for female physicians. Male physicians generally work more hours than female physicians, but the differential decreases with age.

Alternate Measures of Physician Supply

Some researchers have speculated that because female physicians work fewer hours per week than male physicians and because the proportion of female physicians is increasing, projections of the total physician population may overestimate effective physician supply. We investigated this issue by estimating the *aggregate time worked by patient care physicians*, measured in terms of full-time-equivalent physicians. This measure includes only physicians in direct patient care and weights each physician with the amount of time spent in professional activities. This adjusted measure of physician supply takes into account the changing composition of the physician population. We compared this measure of the effective physician supply with the simple number of physicians.

We considered the population of post-resident physicians—physicians who have completed specialty training—because we wanted to analyze how the changing composition of physician supply will affect the specialty distribution of physician services. Our projection estimates did not indicate how many physicians have completed residency training, so we estimated the future number of post-resident physicians based on the available information—the projected age, sex, country of graduation, and specialty distributions of the future physician population. We used a logistic regression analysis of the 1986 AMA Physician Masterfile data to estimate $q_{i,j,k,l}$, the proportion of physicians of age i, sex j, specialty k, and country of graduation l who have completed residency training. (Analyses available upon request to author.) Assuming that these proportions remain constant throughout the projection, we estimated P, the projected size of the active post-resident physician population, as follows:

$$P = \sum P_{i,j,k,l}$$

$$= \sum [q_{i,j,k,l} \, A_{i,j,k,l}]$$

where $A_{i,j,k,l}$ represents the projected number of active physicians in the corresponding age-sex-specialty-country of graduation category.

We estimated F, the future number of full-time equivalent patient care physicians, as follows:

$$F = \sum F_{i,j,k,l}$$

$$= \sum \left[c_{i,j,k,l} \, h_{i,j,k,l} \, P_{i,j,k,l} \, \frac{1}{58.0} \right]$$

where $c_{i,j,k,l}$ is the estimated proportion of physicians in each sex-age-specialty-country of graduation category who are in patient care, derived from our logistic regression analysis of the 1986 Masterfile data; $h_{i,j,k,l}$ is the estimated number of hours worked by patient care physicians, derived from the ordinary least squares analysis of the SMS data; and 58.0 is the average hours per week worked by patient care physicians in 1987.

Table 4 presents historical and projected data for the number of active post-residency physicians, and the number of full-time-equivalent post-residency patient care physicians. Both measures show that physician supply is growing rapidly. Projected growth is slightly slower for the full-time-equivalent measure.

The bottom line of Table 5 presents the number of full-time-equivalent patient care post-residency physicians divided by the total number of post-residency physicians. This ratio describes how effectively the projected physician population will provide time to patient care in terms of 1986 full-time-equivalent patient care physicians. During the 1970–86 historical period, the ratio increased slightly from .888 to .896. Despite the growing proportion of female physicians throughout this time period, the average physician provided more patient care because: the proportion of patient care physicians increased; and a greater proportion of physicians were entering the middle parts of their careers, during which they provide the most patient care.

Between 1986 and 2010, the ratio is projected to decrease from .896 to .854, representing more than a 4 percent

TABLE 4—The Number of Active Post-residency Physicians and the Number of Full-Time-Equivalent Post-residency Patient Care Physicians by Age, Specialty, and Country of Graduation, for Years 1970, 1986, 2000, and 2010

	Historical		Projected	
	1970	1986	2000	2010
Active Post-residency Physicians	262,694	438,935	567,300	611,100
Full-Time-Equivalent Post-residency Patient Care Physicians	233,400	393,500	495,700	522,100
Ratio of Full-Time-Equivalent Post-residency Patient Care Physicians to Active Post-residency Physicians	.888	.896	.873	.854

SOURCE: AMA Physician Masterfile, 1970–86, Department of Data Release Services, Division of Survey and Data Resources; AMA Demographic Model of the Physician Population, AMA Center for Health Policy Research, 1988.

TABLE 5—Percent Change in the Number of Active Post-residency Physicians and the Number of Full-Time-Equivalent Post-residency Patient Care Physicians, 1986–2000 and 1986–2010

	Percent Change in Post-residency Population			
	1986–2000		1986–2010	
	All Active	Full-Time-Equivalent Patient Care	All Active	Full-Time-Equivalent Patient Care
Total	29.2	26.0	39.2	32.7
Sex				
Male	15.6	13.8	15.1	11.5
Female	120.0	120.7	199.3	197.4
Country of Graduation				
USMG	34.7	32.0	49.5	43.6
FMG	11.6	6.8	5.9	−2.1
USFMG	64.3	62.1	92.4	85.5
Foreign-born FMG	3.7	−1.7	−7.1	−15.7
Age (years)				
< 36	−19.5	−21.3	−21.1	−23.0
36–45	16.0	13.8	2.6	−1.3
46–55	76.0	73.0	71.8	67.1
56–65	30.6	30.0	100.2	97.2
> 65	71.2	67.9	121.1	115.5
Specialty				
General/Family Practice	11.8	10.7	17.0	14.4
General Internal Medicine	54.1	51.1	72.5	64.6
Medical Subspecialties	38.4	32.7	58.7	47.4
General Surgery	25.7	24.0	25.5	21.5
Surgical Subspecialties	18.4	15.2	21.2	15.8
Pediatrics	47.1	41.3	67.9	56.5
Obstetrics/Gynecology	26.6	21.2	37.6	28.4
Radiology	30.9	26.3	41.5	33.9
Psychiatry	13.8	8.9	15.7	7.8
Anesthesiology	41.5	37.9	58.9	51.9
Pathology	14.4	8.5	12.0	3.2
Emergency Medicine	57.6	53.0	87.4	77.4
Other Specialties	20.0	15.8	22.4	14.4

SOURCE: AMA Physician Masterfile, 1970–86, Department of Data Release Services, Division of Survey and Data Resources; AMA Demographic Model of the Physician Population, AMA Center for Health Policy Research, 1988.

decrease in the amount of patient care provided by the average physician. This finding is attributable to the growing proportion of women physicians who, on average, work fewer hours per week and are somewhat less likely to be in patient care; and to the growing proportion of older physicians, who work fewer hours per week.

Table 5 shows the percent growth in the two supply measures for the various components of the physician population in 1970–86 and 1986–2010. Both measures indicate that the physician population is growing rapidly, although the growth rate is gradually diminishing. Both indicate that: the percent growth is substantially larger for female physicians than for male physicians; growth rates for emergency medicine, general internal medicine, and pediatrics are greater than for other specialties; and the number of physicians over age 65 will grow faster than the physician population as a whole.

More important for this analysis are the subtle differences between the two measures attributable to changes in the composition of the physician population. In both time periods, the percent increase of active post-residency physicians is several percentage points greater than their full-time equivalents, i.e., the growth of the physician population is slightly faster than the growth of effective physician supply.

For almost all the components of the physician population, the percent increase in active physicians exceeds the percent increase in full-time-equivalent patient care physicians. Female physicians are an important exception. Between 1986 and 2000, disproportionate numbers of female physicians will enter the middle age categories when they spend the most time in patient care. On the other hand, male physicians will be entering the older age categories and thus are expected to work fewer hours per week. Due to this difference, the sex differential in the rate at which physicians provide patient care will decrease in the future.

Discussion

If current patterns persist, the rapidly growing proportion of female physicians will affect both the amount and type of health care available in the future. We have presented data on several important past differences between male and female physicians. The projected growth of the physician population slightly overstates projected growth in the effective supply of physician hours worked. However, the actual magnitude of this difference is small, especially when compared to the rapid growth projected for the physician population.

It is quite likely that, as the number of female physicians increases over time, their practices will more closely resemble those of male physicians. Several factors suggest that some of the differences between male and female physicians may be declining. The specialty distributions of male and female physicians appear to be becoming more similar, especially among the younger age groups. Female physicians are no longer disproportionately represented among foreign-born FMGs. Even though female physicians have higher retirement rates than male physicians, their expected work lives are just as long as male physicians due to their lower mortality rates. As career patterns of male and female physicians become more similar, the classic measures of physician supply will be more reliable indicators of long-term growth.

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New Quarterly Periodical: *Accountability in Research*

A new quarterly journal, entitled *Accountability in Research: Policies and Quality Assurance*, was launched in September 1989 to serve as a forum for critical examination of issues involving the integrity of scientific data. The papers appearing in the first issue are based, in part, on presentations from the First International Conference on Scientific Data Audit, Policies, and Quality Assurance held May 8–10, 1988 at the University of Maryland, College Park. Adil E. Shamoo, University of Maryland School of Medicine, has been named Editor-in-Chief of the new journal.

According to "A Note from the Editor" in the inaugural issue, "Researchers today perceive a growing problem of error and fraud in the vast research and development enterprise. The growing complexity of research, the increased distance between those who gather, manipulate, and evaluate data, and the mounting financial and professional pressures on researchers all contribute to this problem. In recent years, the need for a broad-based consensus on ways of ensuring the integrity of scientific data has become increasingly clear," Shamoo wrote.

The journal will speak to a broad, interdisciplinary range of scientists and administrators in academia, industry, and government. "While relevant to many disciplines," he said, "it will focus in particular on biotechnology, chemistry, medicine, pharmacology, toxicology, pathology, and environmental science. Topics to be addressed include: historical perspectives on data integrity; auditing data as a management tool; research policies and their impact on science and the public; ethical and legal issues; relevant data analysis methods and technology; and legislative and regulatory aspects."

The journal welcomes divergent views on these and related topics from research scientists as well as those in the fields of accounting, law, economics, statistics, management studies, public policy, and information science.

Submissions to the journal should be sent in triplicate to the Editor-in-Chief, Adil E. Shamoo, Department of Biological Chemistry, University of Maryland School of Medicine, 660 West Redwood Street, Baltimore, MD 21201.

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Cryptosporidium Infection in Oregon Public Health Clinic Patients 1985–88: The Value of Statewide Laboratory Surveillance

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Abstract: To evaluate the utility of statewide laboratory *Cryptosporidium* surveillance, we screened stools from all 5,256 patients evaluated at local health departments for parasitic disease from January 1985 through June 1988. Fifty-seven patients (1.1 percent) were found to have *Cryptosporidium*. Seasonal peaks in positivity were observed in the spring, summer, and early autumn months. In children, younger age was associated with higher positivity rate of cryptosporidiosis. As a result of these surveillance efforts, Oregon's first known outbreak of cryptosporidiosis was detected and investigated during 1988. Twenty-five persons were infected, including children, parents, and staff associated with two day care centers. The cost of routine screening for *Cryptosporidium* was $1.13 per specimen in our laboratory, and we consider it useful. (*Am J Public Health* 1990; 80:305–308.)

Introduction

With reports from at least 26 countries,[1-4] *Cryptosporidium* is now well recognized as a pathogen in immunocompetent patients seen for diarrheal disease in a variety of clinical settings. In the United States, most cases are derived from a continuous human reservoir sustained by fecal-oral transmission, with occasional zoonotic infections acquired from cattle, goats, lambs, and domestic animals.[5,6] Surface water may also be an important source of infection.[7,8] Several outbreaks have been reported in day care centers.[9-11] The prevalence of human cryptosporidiosis differs greatly between geographic areas, even within the United States,[12,13] and there is also some evidence for seasonal fluctuation.[14-16]

Many laboratories are currently considering the addition of *Cryptosporidium* screening to the battery of routine tests used for all stools submitted for parasitology. Although *Cryptosporidium* screening is relatively simple, it requires unique staining and examination apart from the usual methods, and its implementation raises issues of cost and work flow which must be addressed. Its usefulness depends upon several factors, including the local prevalence of the agent, the potential for transmission in the population screened, and the cost of screening. There is little information available to assist laboratories and disease control programs in deciding whether to adopt *Cryptosporidium* testing as a part of their routine parasitology protocol.

We have been screening for *Cryptosporidium* infection in immunocompetent public health clinic patients since 1984, and have previously reported our findings in 1,710 patients from our initial year of study.[12] In that report, we noted that although Oregon had not yet experienced the cryptosporidial positivity rate observed in some other areas, we did plan to continue active surveillance. Our ongoing screening during the ensuing three-and-one-half years has allowed us to define further the positivity rate in a selected population, and to detect the first known outbreak of cryptosporidiosis in Oregon. In addition, we have continued to examine the putative association between *Giardia* and *Cryptosporidium* which has been reported by others.[17,18]

Methods

From January 1, 1985 through June 30, 1988, all stool specimens submitted to the Oregon State Public Health Laboratory for parasitological examination from local health department clinics were screened for *Cryptosporidium* in addition to *Giardia* and other parasites. These specimens, which were collected from persons evaluated for parasitic gastroenteritis, were submitted from all of Oregon's local health department outpatient clinics.

All stools were received in 10 percent formalin and concentrated by the formalin-ethyl acetate method.[19] Iodine stained and unstained preparations were examined at both low (100×) and high (400×) magnification for parasitic forms other than *Cryptosporidium*. A smear from each sample was then stained and examined for *Cryptosporidium* in the manner which we have described previously.[12] Briefly, each smear was stained by the auramine-O fluorochrome method and scanned at 400× magnification using a fluorescence microscope with a 390-490 nm light source. Smears were considered presumptively positive if at least one fluorescing object of correct size (and any morphology) was observed per field, or if fluorescing objects morphologically typical of *Cryptosporidium* were observed in any number. *Cryptosporidium* oocysts were seen as typical rings or crescent forms with a mean diameter of 4.5 μm and a resemblance to crenated erythrocytes. Each presumptively positive smear was then overstained with our modification of the Ziehl-Neelsen (Z-N) method for confirmation. The combination of formalin-ethyl acetate concentration plus acid-fast staining has been found to be equivalent to the Sheather sucrose method for recovery of cryptosporidial oocysts.[20] After Z-N staining, the sample was examined at 1,000× magnification and considered confirmed positive if acid-fast oocysts with typical morphology were observed. Positive and negative controls were included in each run and examined in the same manner as clinical specimens. Proper control reactions were a prerequisite for accepting and reporting the concurrent patient results; these controls were correct in 100 percent of the test runs.

Results

Laboratory Surveillance

A total of 5,256 patients (8,396 specimens) were screened for *Cryptosporidium* and other parasites; the percentage who

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had diarrhea was unknown to us. The patients were drawn from a variety of ethnic backgrounds and locales throughout Oregon. Fifteen percent (786) had Southeast Asian surnames; none of these patients were positive for *Cryptosporidium*. Only one of the positive patients was known to have human immunodeficiency virus infection. The age distribution of patients is shown in Table 1.

Fifty-seven of the patients screened (1.1 percent) had stools positive for *Cryptosporidium*, although 25 of these patients were related to the day care-associated outbreak in 1988. Eight of 666 patients with *Giardia* (1.2 percent) had *Cryptosporidium*, compared with 49 of 4,590 patients (1.1 percent) who did not have *Giardia*.

Excluding specimens from the patients who were tested as part of our 1988 outbreak investigation (25 positive and eight negative), the three-and-one-half year statewide "background" *Cryptosporidium* positivity rate among those persons screened was 32 of 5,223, or 0.61 percent. The proportion positive for each age group is given in Table 1. When we compared proportion positive by one-year age increments, we found that among patients less than 18 years old who were tested, younger children were more likely to be infected with *Cryptosporidium*; however, the numbers were small and this difference did not reach statistical significance (p = .07, chi-square analysis for trend).

Cryptosporidiosis Outbreak, 1988

As a result of laboratory surveillance, Oregon's first known community outbreak of cryptosporidiosis that occurred during a 10-week period from mid-April through June 1988 was discovered and investigated. A total of 24 cases and one asymptomatic infection were identified in two day care centers in one community in western Oregon. Cases were defined as having a diarrheal illness with laboratory confirmation of the presence of *Cryptosporidium* oocysts in stool specimens. The epidemic curves by day care center for this outbreak are consistent with initial person-to-person transmission of infection, perhaps followed by a common source outbreak (Figure 1).

The outbreak was brought to the attention of the local health department in the following manner. A child attending Day Care B was found to be positive for *Giardia* and negative for *Cryptosporidium*; however, an investigation revealed that the child had two infant siblings attending Day Care A. One of the two siblings had a stool sample positive for both *Cryptosporidium* and *Giardia*, and the stool sample from the other was negative for both agents. The finding by the laboratory of *Cryptosporidium*, which had not been suspected in this outbreak, alerted the local health department to the need to conduct an epidemiologic investigation centered on this organism.

TABLE 1—*Cryptosporidium* Positivity Rate in Oregon Public Health Patients by Age, 1985–88

Age (years)	No. Patients Examined	% Positive		
		Day Care Outbreak*	Other	Total
<1	188	66.7 (2/3)	1.1 (2/185)	2.1
1–4	1,248	76.5 (13/17)	1.3 (16/1,231)	2.3
5–9	671	100.0 (2/2)	0.6 (4/669)	0.9
10–14	273	100.0 (2/2)	1.1 (3/271)	1.8
15–17	139	0 (0/0)	0 (0/139)	0
≥18	2,540	66.7 (6/9)	0.2 (6/2,531)	0.5
Unknown	197	0 (0/0)	0.5 (1/197)	0.5
Total	5,256	75.8 (25/33)	0.6 (32/5,223)	1.1

*Outbreak April–June, 1988.

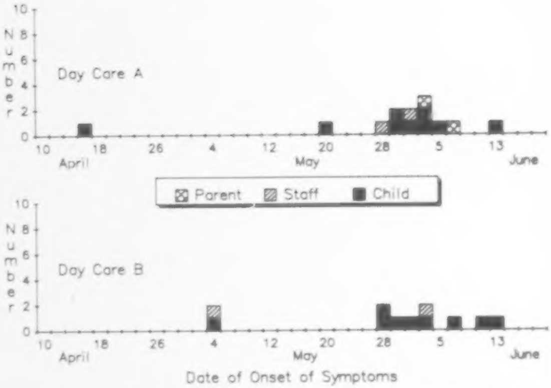


FIGURE 1—Epidemic Curves for Oregon Cryptosporidiosis Outbreak, 1988

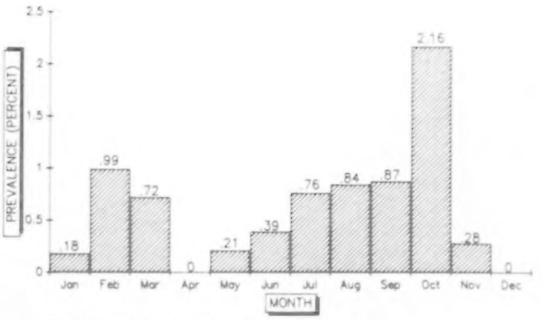
The cryptosporidiosis attack rate for attendees at Day Care A was 90 percent (9/10 children attendees). The onset of the initial case occurred in mid-April. The average age of the children-cases was 1.8 years with a range from 0.8 to 6.0 years. One of the nine case-children also had *Giardia* on stool examination. Both adult staff at Day Care A were infected with *Cryptosporidium*. In addition, two parents of attendees were reported as cases.

At Day Care B the cryptosporidiosis attack rate for attendees was 60 percent (9/15 children). The date of onset of the initial case was May 5, 1988. One attendee had a positive stool specimen, but was asymptomatic, and therefore was not counted as an outbreak-related case. The average age of the case-children was 3.7 years with a range of 2.5 to 6.0 years. Two of the nine case-children also had *Giardia* on stool examination, as did one of the six children who were negative for *Cryptosporidium*. This child was also infected with *Entamoeba histolytica* and *Blastocystis hominis*. In addition to these attendee cases, two of six adult staff were cases. None of the parents were reported as cases.

Excluding the three cases with concurrent giardiasis, who received antimicrobic therapy, all other cases had self-limiting illnesses without treatment; however, the clinical course of one adult case was complicated by a 12-pound weight loss. Among the 10 cases for whom information was available, the average duration of symptoms was eight days with a range of one to 28 days. No one required hospitalization. Control measures included strict handwashing technique and environmental sanitation measures such as cleaning of toys; however, cohorting was not implemented. During the first week of July 1988, repeat stool examinations of all children attendees and infected adults were negative for cryptosporidiosis.

Seasonal Fluctuations in Positivity Rate

When our data were aggregated by month for January 1985 through June 1988, with the 1988 outbreak patients excluded, peak positivity rates were observed in February/March and in summer and early autumn (Figure 2). Given the small number of positive stools in each year, it was necessary to combine the data in this way to identify seasonal patterns.



* Excludes outbreak-related patients, 1988

FIGURE 2—*Cryptosporidium* Prevalence in 6,223 Oregon Public Health Patients Screened for Ova and Parasites by Month, January 1985–June 1988*

Discussion

In Oregon, our finding of 0.61 percent "background" positivity rate in patients screened at all local health department clinics during 1985–88 is consistent with our previous report of 0.35 percent for the same population during 1984.[12] Combining all data for both time periods, including the 1988 outbreak, we have screened 6,966 patients and found 63 (0.90 percent) to be positive for *Cryptosporidium* during four-and-one-half years. Although positivity rates of *Cryptosporidium* vary widely by geography, setting, and population, our results are similar to those reported from Colorado, Texas, Ohio, and British Columbia.[6,13,16,21]

Young children had the highest positivity rate of cryptosporidial infection in our study, even when patients from the 1988 day care-associated outbreak were excluded from the analysis. This finding is consistent with those of other workers and is generally attributed to either greater susceptibility in children or to more efficient fecal-oral transmission among infants and toddlers.[5,16] Less obvious is the source of cryptosporidiosis in older children and adults, many of whom present as sporadic cases without a known source of infection. Many may acquire their infection through contact with young children, as has been suggested in other outbreak situations involving both adults and children.[5,7,11] However, the importance of adult-to-adult transmission (in the immunocompetent) and environmental acquisition is unknown.

Although our findings did not reveal a "statistically significant" association between *Giardia* and *Cryptosporidium* in the population screened or among the day care outbreak cases, co-transmission of these two agents and others may be common during outbreaks. Information is lacking regarding the sensitivity of a single stool examination for the detection of coinfection by these agents. Thus, because of intermittent cyst shedding, differences in incubation period between the two protozoa, the timing of any given sample, and differences in the sensitivity of the laboratory methods, it is possible for one of the pathogens to go undetected, even in a dual infection. For example, some patients excrete atypical *Cryptosporidium* oocysts which are quite fragile and sometimes cannot be stained by the usual methods.[22] Seasonal fluctuations in cryptosporidiosis have also been reported from other countries. Peak prevalences during the late summer and early autumn months have been reported from England[23,24] and Australia;[25] two other studies from the United Kingdom showed peaks during the spring months.[14,15] Among all the seasonal cryptosporidiosis data which are available, those from British Columbia may be the most useful for comparison to our results in Oregon, given the environmental, geographic, and demographic similarities between the two areas. In a one-year study of 7,300 patients with diarrhea in the Vancouver area, the peak incidence was seen in August and September, with over half of the cases occurring in those two months.[16] Our seasonal findings are consistent with those from the Untied Kingdom, Australia, and British Columbia. The interpretation of this seasonality is unclear, but could reflect changing opportunities for acquisition of infection from the environment (e.g., travel, contact with surface water or animals), or for person-to-person transmission.

The relative importance of environmental versus person-to-person transmission of *Cryptosporidium* is still poorly defined, and a better understanding is central to the development of control measures for this agent. Using our experience with *Giardia* as a model, we know that over one-fourth of Oregon giardiasis cases are attributed to drinking surface water, and that there have been several documented community drinking water supply outbreaks of giardiasis in Oregon during the past decade.* There may exist a background of ongoing person-to-person *Giardia* transmission, especially among young children, perhaps augmented by frequent acquisition of the organism from contaminated surface water. Such a pattern may also exist for cryptosporidiosis, but until recently the available technology has limited our ability to examine water for cryptosporidial oocysts.

At least two US waterborne community outbreaks of cryptosporidiosis have been documented, in Texas and Georgia,[26,27] and very recently, the drinking of surface water was identified as an important risk factor in *Cryptosporidium* infections in New Mexico.[7] The Georgia outbreak arose from a sand-filtered and chlorinated community water supply which met all regulatory standards. The presence of cryptosporidial cysts was attributed to infected cattle in the watershed; inadequate flocculation filtration may have allowed the organism to contaminate the drinking water system. Within the past four years, methods have become available for identifying cryptosporidial oocysts in water.[8,28] In Arizona, a study of surface waters during summer 1986 revealed widespread stream contamination with *Cryptosporidium*.[29] A midsummer study of four rivers in Washington and two in northern California showed all six to be contaminated with cryptosporidial oocysts.[8] Similar contamination in nearby watersheds in Oregon seems plausible, especially given the similarities in geography and potential for contamination by dairy cattle and other animals. Comparable studies of Oregon surface water are needed.

An important question about routine cryptosporidial screening of stool samples is whether the cost is justified by the yield. In our laboratory, the direct materials and labor cost for the inclusion of *Cryptosporidium* screening into the usual ova and parasite examination is an additional $1.13 per specimen ($1.03 for personnel, $.10 in materials). The cost-effectiveness of this screening in a given locale will depend upon the individual laboratory cost, the local prevalence of *Cryptosporidium* and the capacity for preventing and controlling outbreaks once infected persons have been identified.

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*Oregon disease surveillance reporting data, unpublished. Portland: Oregon Health Division, 1989.

We conclude that laboratory-based surveillance is useful in defining the *Cryptosporidium* positivity rate and patterns of transmission for a given locale, and in detecting outbreaks when they occur. In areas where *Cryptosporidium* screening is not routinely performed, it is possible tnat the etiology of diarrheal outbreaks, particularly among children, may remain undiscovered or may be incorrectly attributed to other agents. Even in low prevalence populations, laboratory-based surveillance may be justified due to the relatively low cost of screening and the potential for early intervention in outbreaks.

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Call for Abstracts on Dental Public Health Topics

The American Association of Public Health Dentistry has issued a call for abstracts of papers on a broad range of dental public health topics for presentation at its 53rd annual meeting, to be held in Boston, Massachusetts October 10–13, 1990. The theme for the meeting is "Health for All: Opportunities and Responsibilities for Dental Public Health."

Major topics being sought include: access to services, trends in oral diseases or conditions, preventive and educational measures, national/state/local dental programs, delivery systems, infection control, ethics, and impact of state practice acts.

The *deadline for submission of abstracts is April 1, 1990.* An abstract form and information can be found in the November 1989 issue of *Communique* (AAPHD's newsletter) and the Winter 1990 issue of the *Journal of Public Health Dentistry*. Send abstracts to: Alice M. Horowitz, RDH, MA, AAPHD Vice President, 6307 Herkos Court, Bethesda, MD 20817.

For further information, contact AAPHD National Office, 10619 Jousting Lane, Richmond, VA 23235. Tel: 804/272-8344.

A Comparison of Alcohol Sales Data with Survey Data on Self-Reported Alcohol Use in 21 States

PERRY F. SMITH, MD, PATRICK L. REMINGTON, MD, MPH, DAVID F. WILLIAMSON, PHD, AND ROBERT F. ANDA, MD

Abstract: We used data from 21 states that participated in the 1985 Behavioral Risk Factor Surveillance System to compare state-specific per capita self-reported alcohol consumption and the prevalence of three drinking behaviors with state-specific per capita sales. The correlation coefficient for per capita sales and per capita self-reported consumption for the 21 states was 0.81. Per capita sales were also significantly correlated with the prevalence of self-reported heavier drinking, binge drinking, and drinking and driving; the corresponding correlation coefficients were 0.74, 0.59, and 0.51. These findings suggest that states with higher per capita sales of alcohol also have higher rates of self-reported consumption and drinking patterns suggestive of high-risk behavior. (*Am J Public Health* 1990; 80:309–312.)

Introduction

The social, medical, and economic costs of alcohol abuse are a major public health problem in the United States.[1] At the state level, public health officials are paying increased attention to the consequences of problem drinking. Recently, investigators estimated that problem drinking in Minnesota accounted for 1,110 deaths and for $1.4 to $2.1 billion in medical and social expenditures during 1983.[2]

Because direct measures of problem drinking are not always available, indirect measures, such as total sales, are often used to estimate the total alcohol exposure for state populations. For example, one of the 1990 health objectives for the nation is to limit per capita consumption of alcohol to 2.71 gallons per year, an amount estimated from alcohol sales.[3] Sales data at the state level may mismeasure actual per capita consumption because of out-of-state visitors' purchase of alcohol and wastage or stockpiling of alcohol.[4] More importantly, per capita sales may not reflect rates of problem drinking in a population.[5]

A somewhat more direct method for estimating alcohol exposure at the state level is the use of survey data of self-reported alcohol consumption by individuals. However, survey data may not accurately reflect rates of problem drinking for various reasons, including biased self-reporting and problems with sampling.[4,6–8]

Although public health officials use alcohol sales as a measure of typical alcohol consumption and of the frequency of problem drinking, it has not been shown that states with higher alcohol sales have higher levels of per capita consumption and/or higher frequency of drinking behaviors that increase a person's risk for alcohol-related morbidity and mortality. A strong correlation between sales and survey data at the state level would support the use of sales information for monitoring alcohol consumption. To determine the strength of this correlation, we compared state-specific per capita alcohol sales with rates of per capita consumption and three drinking behaviors as reported in surveys administered in 21 states in 1985.

Methods

We obtained state-specific estimates of per capita alcohol sales for 1985 from the Alcohol Epidemiologic Data System of the National Institute on Alcohol Abuse and Alcoholism (NIAAA).[9] These data are based on beverage sales and tax receipt figures for states in which these data are available (including 13 of the 21 states in our study). For the remaining states, data are used on product shipments from major alcoholic beverage industry sources (the Distilled Spirits Council of the United States, United States Brewers Association, and the Wine Institute). For this report, we refer to both sources of data as "alcohol sales." Per capita alcohol sales are determined by dividing each state's total sales in gallons of absolute alcohol for beer, wine, and liquor by the number of state residents aged 14 years and older.

To estimate state-specific rates of alcohol consumption and drinking behaviors based on self-reports, we used 1985 data from the Behavioral Risk Factor Surveillance System. This system uses random-digit telephone surveys for the state-level surveillance of behaviors related to the leading causes of death among adults in the United States. Described in detail elsewhere,[10] the methodology is briefly summarized here.

In 1985, 21 states participated in the surveillance system. Adult respondents were randomly selected from civilian residents, aged 18 years or older, who had telephones. The sample was chosen by a multistage cluster design procedure based on the Waksberg method.[11] Trained interviewers in each state conducted interviews by telephone; calls were made during the day, at night, and on weekends and lasted about 15 minutes each.

In each state, about 100 interviews were completed each month; the annual total per state ranged from 628 to 2,386 (median = 1,177) (Table 1). Response rates were calculated as the percentage of eligible respondents contacted who completed interviews, and ranged from 61–92 percent (median = 83 percent), among the 19 states for which this information was available.

Respondents were asked to report their beer, wine, and liquor consumption, including the number of days in the prior month that they had consumed each type of beverage and the average number of drinks they had consumed per day. They were also asked two questions: "Considering all types of alcoholic beverages, that is, beer, wine, and liquor, as drinks, how many times during the past month did you have five or more drinks on an occasion?" and "During the past month, how many times have you driven when you've had *perhaps* too much to drink?" Those who did not report consuming any alcohol in the month before the interview were considered nondrinkers. For the 21 states, the percentage of nondrinkers ranged from twenty-six to sixty-nine percent (median = forty-one percent) (Table 1).

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TABLE 1—Number of Completed Interviews, Response Rates, and Percentage Nondrinkers, by State, Behavioral Risk Factor Surveillance System, 1985

State	No. of Completed Interviews	% Response Rate*	Percentage Nondrinkers**
Arizona	1,175	76.5	41.0
California	1,372	82.6	37.3
Connecticut	983	***	30.3
Florida	818	61.2	40.4
Georgia	818	81.8	53.8
Idaho	1,179	79.3	47.8
Illinois	1,148	72.8	41.1
Indiana	1,182	90.9	49.2
Kentucky	803	86.6	62.0
Minnesota	2,386	75.2	36.3
Montana	1,183	***	35.6
New York	1,183	79.1	32.8
North Carolina	1,528	89.9	56.7
North Dakota	628	90.4	37.6
Ohio	1,156	75.7	43.8
Rhode Island	1,277	75.4	38.1
South Carolina	1,216	83.7	58.2
Tennessee	1,207	83.8	66.2
Utah	1,162	92.3	69.3
West Virginia	1,177	89.8	63.5
Wisconsin	965	91.4	25.3

*Percentage of eligible respondents contacted who completed interviews.

**Percentage of respondents who did not report consuming any alcohol in the month before the interview. These percentages are adjusted to compensate for the multistage cluster design of the survey and to be representative of each state's population (based on the age and sex distribution of each state's population in 1985).

***Data necessary for computing response rates not available.

To calculate per capita consumption based on self-reports, we assumed that one beer contained 12 oz of 4.5 percent alcohol; one glass of wine 4 oz of 12.9 percent alcohol; and one liquor drink 1 oz of 41.1 percent alcohol. These alcohol percentages are the same coefficients used by the NIAAA to convert the volume of beer, wine and liquor into the volume of absolute alcohol.[9] We added each respondent's estimate of beer, wine, and liquor consumption and multiplied the sum by 12 to determine the total alcohol consumption for each respondent per year. To calculate the mean alcohol consumption by state, we weighted the value of each respondent's total consumption according to his or her probability of selection as a result of the multistage cluster design of the survey.[12] This weighting compensated for the number of persons 18 years of age and older in each household, the number of telephone numbers reaching each household, and variation in cluster size. We also weighted each respondent's total consumption according to the age and sex distribution of the 1985 adult population for each participating state. This procedure provided population-based estimates of per capita consumption that are representative of each participating state.

Using the survey data, we also determined the state-specific prevalence rates of three drinking behaviors: *heavier drinking* (consuming 60 or more drinks during the prior month), *binge drinking* (consuming five or more drinks on at least one occasion during the prior month), and *drinking and driving* (driving at least once during the prior month after having had *perhaps* too much to drink). As was done with the rates of per capita consumption, the state-specific prevalence rates of these behaviors were adjusted to the age and sex distribution of the 1985 adult population for each state, after weighting the responses to account for the multistage cluster design.

We used linear regression to assess the strength of the association between the state-specific per capita alcohol sales and the measures of self-reported alcohol use; the dependent variables were the measures of self-reported alcohol use and the independent variable was per capita sales.

We used computer software from SAS Institute Inc,[13] and Lotus Development Corporation[14] to perform all calculations.

Results

Among the 21 states, per capita alcohol sales varied twofold, from 1.58 gallons per year in Utah to 3.19 gallons per year in Wisconsin (median = 2.49 gallons) (Table 2). Per capita alcohol consumption varied two and one half-fold, from 0.40 gallons per year in Utah to 1.06 gallons per year in New York (median = 0.72 gallons). Self-reported per capita consumption, when expressed as a percentage of per capita sales, ranged from 21.6 percent in Tennessee to 40.5 percent in New York (median = 28.2 percent) (data not shown).

States that ranked higher in alcohol sales also tended to rank higher in self-reported consumption (Table 2). For example, of the four highest-ranking states in sales (Wisconsin, Arizona, California, and Florida), Wisconsin and Florida both ranked second, Arizona ranked sixth, and California ranked seventh in self-reported consumption. The four states that ranked lowest in sales (Utah, West Virginia, Kentucky, and Tennessee) also ranked lowest in self-reported consumption.

Per capita sales showed a strong linear relation with self-reported consumption among the 21 states (Figure 1). The correlation between sales and consumption was 0.81, with a β coefficient of 0.34 (95 percent confidence interval = 0.23, 0.45) (Table 3). The β coefficient indicates that the average per capita increase for this group of states was 0.34 gallons of self-reported alcohol consumption for each gallon increase in per capita sales of alcohol.

We also assessed the strength of association between sales and self-reported consumption among states in two subgroups: the 13 states for which beverage sales and tax receipt figures were used to determine alcohol sales; and the remaining eight states for which alcohol shipments records were used to determine alcohol sales. The correlation between sales and consumption for states in the first group ($r = 0.70$) was similar to the correlation for states in the second group ($r = 0.74$).

The rates of the three drinking behaviors varied from state to state (Table 2). The prevalence of heavier drinking varied threefold, from 3.8 percent in Utah to 11.5 percent in Illinois (median = 7.1 percent). The prevalence of binge drinking varied almost fivefold, from 5.7 percent in Tennessee to 27.8 percent in Wisconsin (median = 16.6 percent). The prevalence of drinking and driving varied sixfold, from 1.5 percent in Kentucky and Tennessee to 9.4 percent in Wisconsin (median = 3.5 percent).

As was found for self-reported alcohol consumption, the prevalence estimates for each of the three drinking behaviors were correlated with per capita sales (Table 3): heavier drinking $r = 0.74$, binge drinking $r = 0.59$, and drinking and driving $r = 0.51$.

Discussion

Our analysis found that alcohol sales and self-reported drinking behaviors are strongly correlated and provide com-

TABLE 2—Annual per capita Alcohol Sales, Self-Reported Consumption of Alcohol, and Prevalence of Self-Reported Drinking Behaviors, by State, 1985

		Self-Reports			
State	Per Capita Sales (gallons)	Per Capita Consumption (gallons)	Rate of Heavier Drinking %	Rate of Binge Drinking %	Rate of Drinking & Driving %
Wisconsin	3.19 (1)*	1.02 (2)	10.4 (5)	27.8 (1)	9.4 (1)
Arizona	3.12 (2)	0.88 (6)	10.7 (4)	17.9 (10)	5.3 (7)
California	3.12 (2)	0.86 (7)	9.3 (7)	16.6 (11)	3.0 (14)
Florida	3.04 (4)	1.02 (2)	10.8 (3)	16.5 (12)	3.0 (14)
Rhode Island	2.94 (5)	0.78 (9)	7.4 (10)	12.3 (16)	2.8 (16)
Connecticut	2.84 (6)	0.99 (4)	11.2 (2)	19.9 (6)	5.7 (5)
Montana	2.81 (7)	0.76 (10)	6.4 (12)	22.3 (4)	5.4 (6)
Illinois	2.72 (8)	0.99 (4)	11.5 (1)	20.8 (5)	6.2 (4)
Minnesota	2.68 (9)	0.86 (7)	7.6 (9)	23.1 (3)	6.8 (3)
New York	2.62 (10)	1.06 (1)	9.5 (6)	18.5 (9)	4.1 (9)
South Carolina	2.49 (11)	0.57 (16)	6.0 (15)	8.2 (19)	2.3 (19)
North Dakota	2.48 (12)	0.61 (14)	4.6 (20)	23.4 (2)	9.3 (2)
Georgia	2.43 (13)	0.57 (16)	6.1 (14)	14.3 (14)	3.5 (11)
Idaho	2.42 (14)	0.67 (13)	7.1 (11)	15.3 (13)	3.3 (13)
Ohio	2.20 (15)	0.72 (11)	8.6 (8)	19.2 (7)	4.3 (8)
Indiana	2.19 (16)	0.71 (12)	5.8 (16)	19.2 (7)	3.8 (10)
North Carolina	2.13 (17)	0.59 (15)	5.6 (18)	11.6 (18)	3.4 (12)
Tennessee	1.94 (18)	0.42 (20)	5.8 (16)	5.7 (21)	1.5 (20)
Kentucky	1.86 (19)	0.56 (18)	6.3 (13)	7.8 (20)	1.5 (20)
West Virginia	1.66 (20)	0.54 (19)	4.9 (19)	13.2 (15)	2.8 (16)
Utah	1.58 (21)	0.40 (21)	3.8 (21)	11.8 (17)	2.5 (18)

*Numbers in parentheses indicate the rank order (high to low) of the 21 states for each measure of alcohol use. The same ranking number was assigned to states with identical values.

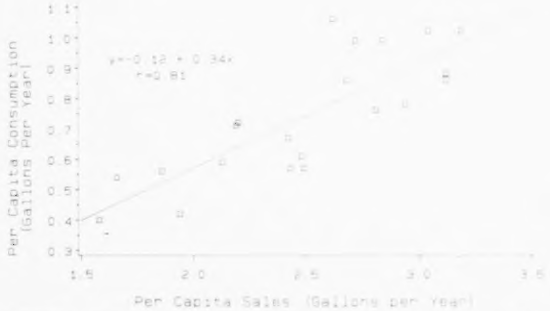

FIGURE 1—Regression Analysis of per capita Sales and per capita Consumption of Alcohol for 21 States, 1985
NOTE: Data points for the 21 states are shown with the fitted regression line, its equation, and the correlation coefficient. Per capita consumption was based on self-reports.

plementary sources of information. However, each measure has its limitations.

State-specific sales may overestimate consumption in states where alcohol is often purchased by out-of-state visitors. An example is the District of Columbia (DC), which participated in the 1985 Behavioral Risk Factor Surveillance System but was excluded from our analysis because DC is more comparable to a large metropolitan area than to a state. Compared with the results for the 21 states, alcohol sales were disproportionately higher in DC relative to self-reported consumption. This finding is due in part to the purchase of alcohol by tourists and residents from nearby states.

An additional limitation to using alcohol sales to monitor alcohol use is that alcohol that is purchased may not be consumed, either because of wastage (unfinished drinks, spillage) or because of stockpiling at the wholesale, retail, and consumer levels. Although the problem of wastage may be negligible,[4] stockpiling can greatly bias estimates of alcohol use.[15]

TABLE 3—Summary of Simple Linear Regression Analyses: Annual per capita Alcohol Sales by Four Measures of Alcohol Use from Self-Reports in 21 States, 1985

Measures of Alcohol Use from Self-Reports	β Coefficient*	r	R^2
Per Capita Alcohol Consumption in Gallons	.34 (.23,.45)**	.81	.66
Prevalence of Heavier Drinking	.03 (.02,.05)	.74	.55
Prevalence of Binge Drinking	.07 (.03,.11)	.59	.35
Prevalence of Drinking and Driving	.02 (.01,.04)	.51	.26

*The β coefficient represents the amount of increase in each measure of self-reported alcohol use for each gallon increase in per capita sales of alcohol.

**Numbers in parentheses indicate the 95 percent confidence intervals for the β coefficients.

In contrast to sales data which provide information only about total alcohol consumption, population surveys can be used to assess both consumption and self-reported drinking behaviors. However, these surveys also have well-recognized problems.[4,6,16] Surveys of self-reported alcohol use uniformly produce lower estimates of total alcohol consumption than do analyses of alcohol sales data.[4,7,8] In previous surveys, self-reported consumption has accounted for only 40–60 percent of alcohol sales, with some reports as low as 26–30 percent of sales.[4] In our study, self-reported consumption accounted for 22–41 percent of alcohol sales. Most telephone surveys are unlikely to include certain segments of the population—such as college students and young men living on military bases—who often are drinkers. Also, population surveys generally undersample heavy drinkers because many of these persons refuse to participate or cannot be located. An additional problem is reporting bias: respondents who are concerned about the social acceptability of their alcohol use may give lower estimates of consumption.

Estimates of self-reported consumption also suffer from the approximations involved in calculating the volume of alcohol consumed per drink. The considerable variation in the size and concentration of alcohol per drink makes estimates of total consumption inexact. This problem has increased with the recent introduction of low-alcohol-content drinks, such as "light" beers and wine "coolers."

Despite all these limitations of sales data and of surveys of self-reported alcohol use, we found a strong correlation between per capita sales and per capita consumption based on self-reports. Per capita sales were also strongly correlated with self-reported heavier drinking, binge drinking, and drinking and driving for the 21 states. These three behaviors do not necessarily identify individual alcohol abusers, nor do they provide accurate measures of problem drinking in a population. However, the prevalence of these behaviors may represent indirect measures of problem drinking by indicating the percentage of a population at increased risk for alcohol-related problems, especially injuries.[17,18]

The results of this study may have been influenced by age considerations. The per capita consumption estimates from sales data were based on each state's number of residents aged 14 and older, as an estimate of the drinking population.[9] Changing the assumed size of the drinking population would affect the calculated rates of per capita consumption from sales data. Furthermore, the surveys included persons only 18 years and older and thus did not reflect consumption and drinking behaviors among younger persons.

Because our study examined ecological associations between alcohol sales and patterns of alcohol use, it was not possible to determine the reasons for the observed correlations between per capita sales and the three drinking behaviors. However, our results are consistent with several theories of alcohol use, including the distribution of consumption model.[19] This theory states that there is a constant relation between per capita alcohol consumption and the prevalence of heavy drinking and suggests that the average alcohol consumption in a population is positively correlated with the prevalence of problem drinking.

We conclude that per capita sales of alcohol generally parallel self-reported consumption and drinking patterns suggestive of high-risk behavior. Although public health officials need to be aware of the limitations of both sales data and self-reports, the high degree of correlation between these two measures of alcohol use suggests that it is reasonable to use alcohol sales for monitoring alcohol consumption when other sources of information are not available.

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Physician Supply Projections to be Re-Examined

Projected manpower needs through the year 2010 in six medical specialties will be re-examined and revised as indicated under a contract awarded to Abt Associates Inc., of Cambridge, Massachusetts. The 13-month, $279,463 project, funded by the Health Resources and Services Administration (HRSA), Division of Medicine, will update projections made in 1980 by the Graduate Medical Education National Advisory Committee (GMENAC) for the fields of family practice, general pediatrics, obstetrics-gynecology, general internal medicine, general surgery, and general and child psychiatry.

The project will involve professional organizations representing these six specialties, and will use Bureau of Health Professions physician supply projections through 2010 adjusted to incorporate the supply of osteopathic physicians. The activity will also include a qualitative assessment of the impact that various economic, social, and behavioral constraints, such as problems of the underserved population, would have on the balance between physician supply and national needs.

For further information, contact Blake Crawford, HRSA, USPHS, DHHS, 5600 Fishers Lane, Room 14-43, Rockville, MD 20857. Tel: 301/443-3376.

Cesarean Section Use and Source of Payment: An Analysis of California Hospital Discharge Abstracts

RANDALL S. STAFFORD, PHD, MS

Abstract: This study assessed the relation between payment source and cesarean section use by analyzing California data on hospital deliveries. Of 461,066 deliveries in 1986, cesarean sections were performed in 24.4 percent. Women with private insurance had the highest cesarean section rates (29.1 percent). Successively lower rates were observed for women covered by non-Kaiser health maintenance organizations (26.8 percent), Medi-Cal (22.9 percent), Kaiser (19.7 percent), self-pay (19.3 percent), and Indigent Services (15.6 percent). Vaginal birth after cesarean (VBAC) occurred more than twice as frequently in women covered by Kaiser (19.9 percent) and Indigent Services (24.8 percent), compared to those with private insurance (8.1 percent). Sizable, although less pronounced, associations between payment source and cesarean section use were noted for the indications of breech presentation, dystocia, and fetal distress. Accounting for maternal age and race/ethnicity did not alter these findings. Variations in the use of cesarean section have a substantial financial impact on health care payors. (*Am J Public Health* 1990; 80:313–315.)

Introduction

Medical care decisions are influenced by non-clinical factors, such as health care financing.[1–3] Cesarean section is the most common hospital surgical procedure in the United States,[4] accounting for 24.4 percent of all deliveries in 1987.* The quadrupling of US cesarean section rates in the past 15 years has led to concerns regarding maternal and perinatal outcome, as well as health care costs.[5]

Past studies suggest that women covered by private insurance have the highest cesarean section rates.[6–8] Women covered by Medicaid, those covered by health maintenance organizations (HMOs) and those without a source of payment (self-pay) have lower C-section rates. Lower C-section rates also are observed for teaching hospitals,[9] public hospitals,[6,8] and salaried physicians.[10] Unfortunately, past research has not controlled adequately for the population characteristics (case-mix) of different payors or organizational settings.

This report investigates the relationship between hospital payment source and the use of cesarean section by employing 1986 discharge abstract data from California hospital deliveries.

Methods

Hospital Discharge Data

The Office of Statewide Health Planning and Development collects information on every non-military hospital discharge in California;[11] data for 1986 included 461,066 hospital deliveries.

Cesarean section use was determined via ICD-9-CM procedure codes. Initial categories for expected source of payment for hospitalization were combined into seven distinct payment methods: Private Insurance (Blue Cross/Blue Shield and other private insurance), Kaiser Permanente, Other (non-Kaiser) HMOs, Medi-Cal (the California Medicaid program), Self-Pay, Medically Indigent Services, and Other Payors. Women with HMO coverage delivering in Kaiser Hospitals were assigned to Kaiser Permanente, all other HMO deliveries were assigned to Other HMOs.

Diagnostic information was available as ICD-9-CM codes. A standard and widely employed hierarchy of mutually exclusive diagnoses which accompany cesarean section was adopted for this study.[12,13] From highest to lowest priority, these were: previous cesarean section (ICD-9-CM 654.2), breech presentation (652.2 and 669.6), dystocia (653 and 660-662, except 661.3), fetal distress (656.3), and all other diagnoses (remaining codes in 642-663, and 669, except 669.1-669.5).[12,13] When more than one diagnosis was present for a delivery, the highest priority diagnoses was assigned regardless of its order on the discharge abstract. Maternal age was defined in five-year categories (<20, 20–24, 25–29, 30–34, 35+). Five race/ethnicity categories were employed: White, Black, Hispanic, Asian and Others/Unknown.

Statistical Methods

Cesarean section rates by payment source for individual primary indications were calculated. Confidence intervals were calculated for the cesarean section rates of each payor. Payors whose confidence intervals do not overlap can be considered different at a level of $p < 0.05$, after accounting for multiple comparisons between payors.[14]

To assess the constancy of the relation between payment source and cesarean section rates, the sample was simultaneously stratified by accompanying diagnoses, maternal age, and race/ethnicity. To further validate the findings derived from the above methods, logistic regression analysis was employed to simultaneously estimate the independent effects of payment source, accompanying diagnoses, maternal age, and race/ethnicity. The results of this analysis closely agreed with those from the tabular analysis (details available upon request to author).

Results

In 1986, 112,730 cesarean sections were performed in California hospitals, a cesarean section rate of 24.4 percent. Private Insurance (35 percent) and Medi-Cal (27 percent) were the dominant payors for hospital deliveries (Table 1).

Cesarean section rates varied widely by payment source. Women covered by Private Insurance had the highest cesarean section rates (29.1 percent) (Table 1). Indigent women had the lowest cesarean section rates, with only 15.6 percent being delivered abdominally.

Address reprint requests to Randall S. Stafford, PhD, Epidemiology Program, School of Public Health, University of California, Berkeley, CA 94720. This paper, submitted to the Journal March 27, 1989, was revised and accepted for publication August 24, 1989.

*Taffel SM, Placek PJ: Current trends in cesarean section in the United States. Presented at the November 1988 meeting of the American Public Health Association, Boston, MA.

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TABLE 1—Cesarean Section Rates by Payment Source, Hospital Deliveries, California, 1986

Payment Source	Total Deliveries	Percentage of all Deliveries	C-Section Rate (95% CI)
Private Insurance*	161,847	35.1	29.1 (28.8, 29.4)
Other HMOs	46,034	10.0	26.8 (26.3, 27.4)
Medi-Cal	122,792	26.6	22.9 (22.6, 23.2)
Kaiser Permanente	51,589	11.1	19.7 (19.2, 20.1)
Self-Pay	49,292	10.7	19.3 (18.8, 19.8)
Indigent Services	10,322	2.2	15.6 (14.6, 16.5)
Other Payors**	19,190	4.2	20.2 (19.5, 21.0)
Total	461,066	100.0	24.4 (24.3, 24.6)

*Includes Blue Cross/Blue Shield and Other Private Insurance.

**Includes Medicare (658 deliveries), Workers' Compensation (104), Title V (208), Other Governmental Payors (13,220), Other Non-governmental Payors (3,709), and No Charge (1,292).

Repeat cesarean sections accounted for 36 percent of all cesareans (Table 2). Large payor differences were observed in women with previous cesarean sections, with an ordering of payors similar to that for all indications (Table 3). Repeat cesarean section rates varied from 91.9 percent for Private Insurance to 75.2 percent for Indigent Services, the rate of vaginal birth after cesarean (VBAC) in indigent women (24.8 percent) was three times that in women with Private Insurance (8.1 percent).

Cesarean section for breech presentation exhibited a distinct pattern by payor. Kaiser had a relatively high breech cesarean rate (88.3 percent); indigent women the lowest (78.5 percent) (Table 4).

Cesareans for dystocia accounted for 32 percent of all cesareans. For this accompanying diagnosis, relatively high cesarean rates were experienced by Private Insurance (65.9 percent, 65.1–66.7), Other HMOs (65.3 percent, 63.7–67.0), Medi-Cal (64.6 percent, 63.5–65.7), and Self-Pay (61.4 percent, 59.7–63.1). Women covered by Indigent Services (56.9 percent, 52.3–61.5) and Kaiser (58.8 percent, 55.4–62.3) had significantly lower rates.

Deliveries with the diagnosis of fetal distress showed a pattern similar to that for all deliveries. Other HMOs (33.0 percent, 30.9–35.0) and Private Insurance (31.5 percent, 30.4–32.6) had the highest rates, while indigent women had the lowest (21.6 percent, 18.4–24.9).

TABLE 2—Cesarean Section Rates by Primary Accompanying Diagnosis, Hospital Deliveries, California, 1986

Primary Accompanying Diagnosis	Deliveries	Percentage of all Deliveries	Cesarean Sections	Percentage of all Cesareans	Cesarean Section Rate
Previous Cesarean	45,188	9.8	40,274	35.7	89.1
Breech	13,926	3.0	12,112	10.7	87.0
Dystocia	55,895	12.1	35,896	31.9	64.2
Fetal Distress	39,058	8.5	11,134	9.9	28.5
Other Diagnoses*	111,156	24.1	12,926	11.5	11.6
None**	195,843	42.5	388	0.3	0.2
Total	461,066	100.0	112,730	100.0	24.4

*Includes the following diagnoses occurring without previous cesarean section, breech presentation, fetal distress, or dystocia: antepartum hemorrhage, hypertension, excessive vomiting, pre-term labor, prolonged pregnancy, other pregnancy complications, maternal infections, other medical conditions in pregnancy, multiple gestation, other malpresentation, anatomical abnormalities of pelvis, fetal abnormalities, placental problems, polyhydramnios, other amniotic complications, other indications for care, umbilical cord complications, and other complications of labor and delivery.

**Includes deliveries lacking any diagnosis classifiable to any of the above categories.

TABLE 3—Cesarean Section Rates for Women with Previous C-Sections, by Payment Source, Hospital Deliveries, California, 1986

Payment Source	Total Deliveries	Percentage of Payor's Deliveries	C-Section Rate (95% CI)
Private Insurance	18,837	11.6	91.9 (91.4, 92.4)
Other HMOs	5,064	11.0	91.7 (90.6, 92.7)
Medi-Cal	11,444	9.3	90.6 (89.9, 91.4)
Kaiser Permanente	4,385	8.5	80.1 (78.5, 81.7)
Self-Pay	3,353	6.8	81.9 (80.1, 83.7)
Indigent Services	660	6.4	75.2 (70.7, 79.6)
Other Payors	1,445	7.5	82.9 (80.3, 89.5)
Total	45,188	9.8	89.1 (88.7, 89.5)

TABLE 4—Cesarean Section Rates for Breech Presentation, by Payment Source, Hospital Deliveries, California, 1986

Payment Source	Total Deliveries	Percentage of Payor's Deliveries	C-Section Rate (95% CI)
Private Insurance	5,510	3.4	89.7 (88.6, 90.8)
Other HMOs	1,381	3.0	89.9 (87.7, 92.1)
Medi-Cal	3,400	2.7	83.9 (82.2, 85.6)
Kaiser Permanente	1,424	2.8	88.3 (86.0, 90.6)
Self-Pay	1,416	2.9	81.6 (78.8, 84.4)
Indigent Services	247	2.4	78.5 (71.4, 85.6)
Other Payors	548	2.9	85.0 (80.9, 89.1)
Total	13,926	3.0	87.0 (86.2, 87.8)

For deliveries with other accompanying diagnoses, women with Private Insurance had cesarean section rates (14.4 percent, 13.9–14.9) nearly double those of indigent women (7.7 percent, 6.3–9.1).

Of the 125 subpopulations formed from all combinations of age-group (five categories), race/ethnicity (five categories), and accompanying diagnoses (five categories), 105 had three or more payors with more than 20 deliveries. Among these subpopulations, Private Insurance had the highest cesarean rates in 42 percent and non-Kaiser HMOs in 30 percent. Medi-Cal (8 percent), Other Payors (8 percent), Self-Pay (6 percent), Kaiser (6 percent), and Indigent Services (1 percent) had the highest cesarean rates less frequently (detailed tables available on request to author).

Discussion

This study indicates that cesarean section use is strongly associated with the source of payment for obstetrical care, after stratifying for available patient characteristics. Differences between payors were particularly large for women with previous cesarean sections, currently a controversial area of obstetrical practice.[15,16]

Because payment source is not randomly assigned to women, stratification by patient characteristics is an important aspect of this study. Two potential limitations in this process must be noted. First, hospital discharge abstracts have a limited number of data elements available, and do not include information on parity, gravidity, and birthweight. Second, potential problems in data accuracy and completeness may diminish the discharge data's ability to reflect clinical differences between patients. In particular, dystocia and fetal distress may be subject to biased recording and are more likely to be recorded if a cesarean section is performed.

Direct or indirect financial incentives, as well as additional factors such as access to technology,[8] physician work schedules,[17] hospital teaching status,[8] hospital ownership status,[6,9] patient socioeconomic status,[18] and medical malpractice concerns may contribute to the observed payor differences. In addition, it is possible that the underlying delivery preferences of women or their physicians may vary systematically by source of hospital payment.

The high cesarean rate for Private Insurance is consistent with the financial incentives of fee-for-service practice that favor higher cesarean section rates. The existence of incentives does not necessarily imply that physicians consciously seek greater income. Such incentives may function via implicit clinical standards.

The moderate Medi-Cal cesarean rates are consistent with incentives to perform cesarean section that may be tempered by the reduced physician and hospital reimbursements provided by Medi-Cal.[19] Women with Medi-Cal also tend to deliver in hospitals where there may be other constraints on cesarean section use.

While HMOs have been shown to have lower cesarean section rates than private insurance,[7,20] the current findings suggest this is the case only in large, hospital-based HMOs. The data indicate that the cesarean section experience of HMOs other than Kaiser closely resembles that of private insurance. Besides Kaiser's incentive to avoid excessive cesarean section use, the prominence of peer review and fixed work schedules also may contribute to the lower Kaiser cesarean rates.

The desire of Self-Pay women to avoid the higher cost of cesarean section may discourage cesarean section use. In addition, Self-Pay women make disproportionate use of public hospitals, where additional institutional constraints may reduce cesarean use.

The low rates of cesarean section in women covered by Indigent Services is consistent with the economic constraints of the county hospitals where these women deliver. At the same time, many of these hospitals have teaching programs which, like Kaiser, are more likely to have fixed work schedules and emphasize peer review.

The potential costs savings to payors of fewer cesarean sections may be gauged using 1986 estimates of physician and hospital costs, and Kaiser's moderate cesarean section rate of 19.7 percent. Insurance claims data for the western US indicate that in 1986 cesarean sections were 84 percent more costly than vaginal deliveries ($5,000 versus $2,720).[21] Based on these figures, if the cesarean section patterns of Kaiser were adopted, private insurance plans, non-Kaiser HMOs and Medi-Cal potentially would save $51 million annually by avoiding 22,500 cesarean sections. This may be an underestimate of cost savings, however, because Kaiser's rate is nearly twice that suggested as an optimal cesarean section rate.[22,23]

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Public Health Briefs

The Effects of Case Definition in Maternal Screening and Reporting Criteria on Rates of Congenital Syphilis

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Abstract: Reports of congenital syphilis in 1987 were reviewed to determine how new national guidelines for defining congenital syphilis would influence reported rates in Los Angeles County. After reviewing all reported reactive serologic tests for syphilis, we found 166 additional cases, resulting in a 426 percent increase in the 1987 reported rate of congenital syphilis. Rates of congenital syphilis are dependent upon surveillance practices, screening policies, and case definition. (*Am J Public Health* 1990; 80:316–317.)

Introduction

Until January 1988, the Centers for Disease Control (CDC) had asked health departments to report cases of congenital syphilis based upon Kaufman's criteria.[1] This complicated system classified infants as "definite," "probable," "possible," and "unlikely" based upon symptoms and laboratory findings at birth and serological tests collected over time. A major shortcoming is the variability with which follow-up data might be available, making the accurate reporting of cases difficult. A second problem was that Kaufman's criteria excluded stillborn infants of mothers with untreated syphilis.[2]

In January 1988, the CDC issued new guidelines for classifying and reporting cases of congenital syphilis.[3] The new guidelines broadened the case definition to include all liveborn and stillborn infants with reactive serologic tests for syphilis (STS) delivered by women with a history of untreated syphilis. Such infants are now considered to be "compatible" with a diagnosis of congenital syphilis, regardless of symptoms. This category essentially combined and simplified the "probable" and "possible" classifications of Kaufman's criteria, as well as broadening it to include fetal deaths.

In order to determine how adoption of the new CDC criteria would affect the reported rates of congenital syphilis, the Los Angeles County (California) Sexually Transmitted Disease Program reviewed its 1987 congenital syphilis caseload.

Methods

In Los Angeles County, all cases of syphilis and all laboratory findings consistent with syphilis must be reported to the Sexually Transmitted Disease (STD) Program of the Department of Health Services.[4] Because California reporting requirements only include symptomatic cases of "probable" and "possible" congenital syphilis, asymptomatic seropositive infants with no other laboratory findings have not been routinely counted in Los Angeles morbidity reports. Instead, these infants have been classified as "epidemiologically treated" or non-cases, because they received treatment for syphilis without supportive evidence that they were truly infected. All reports of "epidemiologically treated" infants were reviewed to distinguish those infants whose mothers had a documented history of adequate treatment prior to delivery ("unlikely" cases) from those infants whose mothers were not treated for syphilis prior to delivery ("compatible" cases). Unreported infants with negative STS born to untreated syphilitic mothers were not included in the review.

To determine the number of fetal deaths in 1987 that occurred among the women with untreated syphilis, we matched all fetal death records (n = 1133) with all STD records by last name and address. As matching in this way might not identify all such outcomes, we reviewed fetal death and syphilis serology records from three public and two private hospitals reporting the largest numbers of congenital syphilis cases in 1987.

Results

Using California's congenital syphilis reporting criteria, Los Angeles County reported 39 cases of congenital syphilis in 1987. An additional 189 infants received presumptive treatment in 1987. Of these, 115 could be classified as "compatible" with a diagnosis of congenital syphilis under the new CDC criteria. In addition, the review of fetal death

TABLE 1—Race-Specific Rates* of Congenital Syphilis in Los Angeles County (California) for 1987

Race	Reportable by Existing Criteria		Reportable by New CDC Criteria		Proportionate Increase
	NUMBER	RATE	NUMBER	RATE	
White	1	2.3	11	25.7	11.1
Black	26	115.3	117	519.0	4.5
Hispanic	12	15.0	59	73.9	4.9
Asian-American	0	0.0	2	16.1	
Unknown	0	0.0	16	391.8	
Total	39	26.0	205	136.7	4.9

*Cases/100,000 total births

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TABLE 2—Compatible Congenital Syphilis by Institution of Birth, Los Angeles County, 1987

	Public hospitals		Private hospitals			
Race	n	Rate per 100,000[a]	n	Rate per 100,000[a]	Relative Risk	95% CI
White	3/959	312.8	8/41,875	19.1	16.37**	(6.17, 43.5)
Black	51/2,919	1,747.2	66/19,624	336.3	5.19*	(3.75, 7.20)
Hispanic	39/26,681	146.2	20/53,149	37.6	3.88*	(2.36, 6.40)
Asian	2/495	404.0	0/11,904	0		
Unknown	4/331	1,208.5	12/3,751	319.9	3.78**	(1.32, 10.80)
TOTAL	99/31,385	315.4	106/130,303	813.5	3.88*	(3.01, 5.00)

[a] per 100,000 births (liveborn and stillborn)
*calculated using Yates Corrected Chi-Square
**calculated using Fischer's Exact Chi-Square

records identified 42 among women with untreated syphilis, four among women with no documentation of treatment for syphilis, and five among women who had been treated within the week preceding the stillbirth. The total of 205 cases represent nearly a five-fold increase in the risk of congenital syphilis in Los Angeles County.

When stratified by race, the 39 officially reported cases in 1987 included only one White and no Asian-Americans (Table 1). When all 205 cases were included all race groups were represented. The new rates of congenital syphilis were highest among Blacks, 7.0 times and 20.2 times higher than Hispanics and Whites, respectively.

When cases were divided by ethnicity and source of report, infants born at public hospitals were more likely to be diagnosed as being "compatible" with congenital syphilis than infants born at private hospitals for all race groups (Table 2).

Discussion

Using the expanded case definition, cases of congenital syphilis increased nearly five-fold in 1987, and even this increase is probably an underestimate. Laboratory reporting of reactive STS results is often variable and, particularly in the private sector, pregnancy status may not be reported with a reactive serology. STS is not mandatory at delivery in California; 48 percent (19/39) of the symptomatic reported cases in 1987, diagnosed only when they developed the florid symptoms of congenital syphilis, were missed at delivery. It is probable that many more at-risk infants were not identified due to lack of screening.

The high rate of congenital syphilis among Blacks is consistent with the rates of reported infectious syphilis in Los Angeles (Blacks 300.6/100,000, Whites 10.3/100,000, Hispanics 42.5/100,000).[5] In our series Black infants were more likely to be among those who were not diagnosed until symptoms developed many weeks after delivery (18/19). They were less likely to have prenatal care and more likely than Hispanics to deliver in private sector hospitals where delivery screening is not routine.

Public hospitals in Los Angeles County routinely perform the STS in the first and third trimesters of pregnancy and at delivery. While women who receive care there may be at greater risk for syphilis, they are more likely to be diagnosed with syphilis than in private hospitals where screening is at physician discretion. Since the distribution of livebirths and stillbirths in public hospitals differs from the distribution countywide, overrepresenting Hispanics and underrepresenting all other race groups, these racial-ethnic comparisons may be inaccurate.

The 205 cases of congenital syphilis defined by the new CDC criteria are a better reflection of the extent of the local congenital syphilis problem than the 39 defined by a modified Kaufman's criteria because more infants at risk are included—not just those few who develop classic symptoms of congenital syphilis. While the new CDC criteria overestimate the cases by including uninfected infants who may have a reactive STS due to passive transfer of maternal antibody, many syphilis-related stillbirths are uncounted because the mother is not screened at delivery and the association is never made. Because there is no way to differentiate whether an asymptomatic infant born to a mother with untreated syphilis is healthy or infected, all at-risk infants are treated and should be counted, regardless of whether they have reactive STS. In Los Angeles County, because four infants born with non-reactive STS in 1988 were not appropriately treated at delivery (their mothers' sera were reactive), they subsequently developed florid symptoms of congenital syphilis.

These data demonstrate that reported cases of syphilis are greatly affected by both definition, reporting criteria, and maternal delivery screening practices. For rates to be comparable between reporting jurisdictions, a clear case definition and universal delivery screening for mothers of both stillbirths and livebirths are necessary.

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End State Renal Disease among Native Americans, 1983–86

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Abstract: We used data reported to Medicare from 1983 through 1986 to determine the incidence of end-stage renal disease (ESRD) among Native Americans and Whites in the United States. The 1,075 Native American cases represented an annual incidence, age-adjusted to the White population, of 269 per million, 2.8 times the rate for Whites. Fifty-six percent of Native American cases and 27 percent of the White cases were attributed to diabetes, indicating that ESRD is a major problem. Diabetes control provides the greatest opportunity for prevention. (*Am J Public Health* 1990; 80:318–319.)

Introduction

Chronic diseases, including end-stage renal disease (ESRD) and diabetes, are major public health problems for Native Americans; in the period 1981–83, their age-adjusted mortality rate due to renal disease was 11.7 per 100,000, while the 1982 rate for all US races was 4.5 (National Center for Health Statistics data provided by Thomas Welty, MD, MPH). Although the incidence of ESRD among all racial groups appears to have risen dramatically since the mid-1970s, when federal funding for ESRD treatment was initiated, the increase has been more rapid among the minority populations than among the White population.[1,2]

To compare the distribution of ESRD treatment onset among Native Americans and Whites, we examined the Health Care Financing Administration's ESRD Medical Information System. This system, designed for fiscal management, includes data on over 90 percent of ESRD treatment in the United States.[1]

Methods

The HCFA Medical Evidence Report form is completed for every patient receiving dialysis or renal transplantation reimbursed by HCFA by the medical care provider. Race categories are "Native American" (American Indian or Alaskan Native), Asian or Pacific Islander, Black, White, and unknown. (Ethnicity has not been uniformly recorded.) We reviewed data for all ESRD referrals who were identified as "Native American/American Indian" and as "White" for the period 1983–86: age, sex, year of treatment initiation, state of residence at the time of initiation, and primary diagnosis.

To determine the extent of one form of race misclassification, the case reports designated as American Indian for the period 1982–86 were examined for any names that suggested apparent origin from India or other Asian or Pacific Island countries.

The estimated number of diabetic Native Americans in 1987 was 72,000 in a total population of 1.6 million (diabetes prevalence from Indian Health Service ambulatory care data, provided by Sarah Valway, DMD, MPH); and the national figure for diabetic Whites in 1985 was 4,892,000 out of a total population of 197,576,000.[3] We used these national estimates to calculate the incidence of ESRD among Native Americans and Whites and to determine the association between ESRD and diabetes. Population data from the 1980 US Census were used to calculate overall, age-specific, and sex-specific rates in Native Americans and Whites. We used the direct method to age-adjust Native American ESRD rates to the 1980 White population.

Results

Of cases of ESRD treatment initiation reported to Medicare from 1983 to 1986, 1,075 were categorized as American Indian and 75,291 as White. Fifty-eight percent of the Native American cases resided in the five states with the largest Native American population, in which the crude incidence was 219 per million, compared to 140 per million in 42 other states reporting 1 to 50 American Indian cases.

The age-adjusted incidence of ESRD for Native Americans was 269 per million, 2.8 times that for Whites (Table 1). The age-specific rate ratios for Native Americans were highest in the 45–64 age group.

The proportion of ESRD attributed to diabetes was greater for Native Americans than for Whites (55.5 percent vs. 26.9 percent, $p < 0.01$) (Figure 1). In the Native American total population, the age-adjusted incidence of ESRD attributed to diabetes was 149 per million, 5.8 times the rate for Whites.

In the estimated Native American diabetic population, the incidence of ESRD attributed to diabetes was 2,073 per million, 2.0 times the rate for Whites. The incidence of ESRD associated with other causes in the non-diabetic population was 78 per million for American Indians, 1.1 times the rate for

TABLE 1—Annual Incidence of ESRD Treatment Reported to Medicare among American Indians and Whites, United States, 1983–86

	Incidence[a]		
Age (years)	American Indian	White (referent)	Rate Ratio (95% CI)
0–14	9	8	1.1 (0.4,3.0)
15–24	31	24	1.3 (0.7,2.4)
25–34	93	58	1.6 (1.0,2.4)
35–44	229	94	2.4 (1.8,3.4)
45–54	546	124	4.4 (3.4,5.7)
55–64	889	220	4.0 (3.2,5.1)
65–74	1042	340	3.1 (2.3,4.0)
75+	657	274	2.4 (1.5,3.8)
Total			
Crude	168	95.3	1.8 (1.6,2.0)
Age-adjusted	269[b]	95.3	2.8 (2.6,3.1)

[a] Rates per million population; based on the 1980 US Census.
[b] Direct age-adjustment using 1980 US Census estimate.

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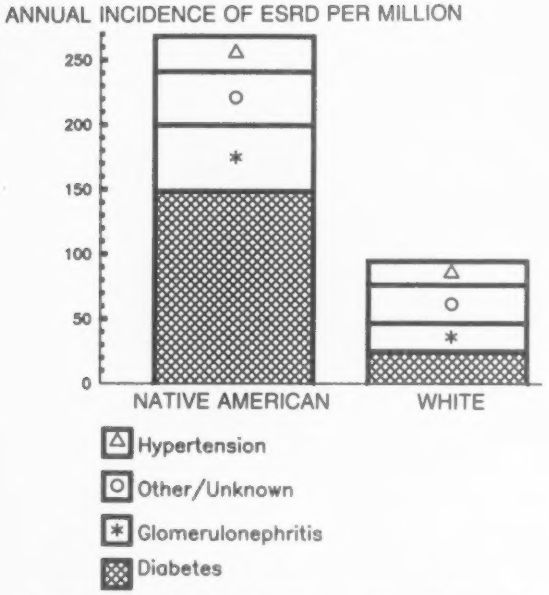

FIGURE 1—Incidence of End-Stage Renal Disease, by Primary Diagnosis, Native Americans and Whites, 1983–86*
***Native American Rate Age-Adjusted to White Population**
SOURCE: ESRD data from Medicare, Health Care Financing Administration

Whites. The relative risk for ESRD associated with diabetes is 27 for Native Americans, compared with 14 for Whites.

Review of the ESRD cases from 1982–86 classified as American Indian revealed that the names in 7.4 percent of the cases (93/1265) were highly suggestive of origin from India or other Asian or Pacific Island countries. This apparent race misclassification was less prevalent in cases attributed to diabetes (3.8 percent or 26/678) than in those attributed to other causes (11.4 percent or 67/587), and in the reports submitted from the five states with more than 50 Native American ESRD cases (4.7 percent or 35/739) than from the other states (11.0 percent or 58/526). There was no association between apparent race misclassification and the sex of the referred patient, the age of the referred patient, the year of referral, the incidence rate of ESRD, or the size of the Native American population within the state of referral.

Discussion

For Native Americans, the age-adjusted incidence of ESRD treatment reported to Medicare from 1983 through 1986 was 2.8 times that for Whites. The majority of Native American ESRD cases were attributed to diabetes; the diabetes-attributed incidence was 5.8 times that for Whites. These data are consistent with other measures of the burden of diabetes for Native Americans.[4] While the prevalence of diabetes is higher in Native Americans, the higher incidence of diabetes-attributed ESRD in its diabetic population than among Whites with diabetes suggests that severity of diabetes and other factors among Native Americans are also involved. However, this finding may also be due to an underestimate of the prevalence of diabetes in Native Americans.

Our review of names of cases designated as American Indian indicates that the observed incidence of ESRD is not substantially inflated by one type of race misclassification, that of apparent Asians as American Indian. Because of mixed heritage, there may also be some misclassification between Whites and American Indians. It is likely that misclassification of American Indians as White exceeds misclassification of Whites as American Indian, suggesting that the observed American Indian incidence may be an underestimate. Further study of race misclassification may add to the utility of the ESRD data set for surveillance by the Indian Health Service.

Although the type of diabetes was not designated, Native Americans have a high prevalence of non-insulin dependent diabetes mellitus and a low prevalence of insulin-dependent diabetes, which is consistent with their older age of onset of ESRD. A recent study of the Pima Indians demonstrated a high incidence of ESRD attributed to non-insulin dependent diabetes, and a strong association with its duration.[5]

These findings indicate the need for prevention of ESRD in Native Americans. Potential interventions include the primary prevention of non-insulin dependent diabetes[6] and early identification of renal disease in persons with diabetes, to intensify diet and blood pressure control.[7]

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Mortality from Infectious Diseases among New Mexico's American Indians, Hispanic Whites, and Other Whites, 1958–87

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Abstract: To examine ethnic differences in infectious disease-related mortality in New Mexico's American Indian, Hispanic White and other White populations, we analyzed vital records data from 1958 to 1987. We found that for most infectious causes, American Indians had the highest mortality rates, followed by Hispanics. The state's minority populations remain at increased risk for infectious disease mortality. (*Am J Public Health* 1990; 80:320–323.)

Introduction

Trends in infectious disease mortality among minority groups in the United States have not been adequately examined. New Mexico has three major ethnic populations: American Indians, Hispanic Whites and other Whites. The cultural and genetic backgrounds of these populations are distinct, and the three ethnic groups have differing levels of education and income, with over 40 percent of American Indians and 23 percent of Hispanics living below the poverty level.[1] Many American Indians in New Mexico live traditional life-styles, often in crowded homes with no running water.[2] Despite gradual economic improvement in the state, many disadvantaged minority persons still live in conditions similar to those in developing countries,[3–5] with the attendant risks of acquiring many types of infectious diseases. To examine ethnic differences in infectious disease-related mortality in New Mexico, we analyzed vital records data collected from 1958 to 1987 for American Indians, Hispanic Whites, and other Whites in the state.

Methods

We obtained coded death certificate data for New Mexico residents for the years 1958–87 from the New Mexico Bureau of Vital Statistics. Cause of death on the certificates was coded according to seventh, eighth, and ninth revisions of the International Classification of Diseases (ICD) for the years 1958–68, 1969–78, and 1978–87.[6–8] Conforming to ICD categories, we examined data for: tuberculosis, other (nontuberculous) infections and parasitic diseases, influenza, pneumonia, kidney infections, and meningitis. Due to ICD classification coding changes, we did not examine diarrheal disease.

The Bureau of Vital Statistics assigned ethnicity according to information on the death certificates. Hispanic ethnicity was determined by surname of the decedent and the decedent's parents, and specific statements on the death certificates; other White ethnicity by coding as "White" on the certificate, and neither a Spanish surname nor other information to indicate otherwise; and American Indian ethnicity by information on the death certificate. American Indian cases were primarily of the Navajo, Pueblo, Zuni, and Apache tribes. Small numbers of cases prevented inclusion of other racial/ethnic groups in the study. The Bureau of Vital Statistics' assignment of Hispanic and other White ethnicity correlated highly with self-reported ethnic status.[9,10] Among Hispanic cases, no attempt was made to distinguish between recent immigrants to the state and the descendants of early settlers to New Mexico.

Denominators for rate calculations were derived from the censuses of 1960,[11–13] 1970,[14,15] and 1980.[1] Since different techniques were used in each census by the US Census Bureau to identify ethnicity, we adjusted estimates of the Hispanic population from 1960 to 1980 to account for the different enumeration procedures.[9,10] Because of coding changes among the ICD revisions, we standardized case counts to conform to the coding scheme of the eighth revision with the National Center for Health Statistics' comparability ratios.[16,17]

Age-specific and age-adjusted mortality rates were calculated by infectious disease category for the five-year periods, 1958–62, 1963–67, 1968–72, 1973–77, 1978–82, and 1983–87. Age-adjusted rates were standardized to the 1970 US standard million population. We used the Statistical Analysis System[18] for our analyses.

Results

The data show substantial differences in infectious disease mortality among New Mexico's three major ethnic populations during the 30-year study period (Tables 1, 2). Age-adjusted mortality rates for most categories of infectious diseases were generally highest among American Indians, followed by Hispanics, for each time period of the study. The greatest rate differences were evident in most periods for tuberculosis, pneumonia, meningitis, other infections, and parasites, with American Indian rates far exceeding rates for Hispanics and other Whites (Tables 1, 2). Ethnic differences for influenza and kidney infections were less striking (Tables 1, 2).

Dramatic declines in mortality rates were observed for American Indians for tuberculosis and for other infections and parasites (Tables 1, 2). However, mortality for tuberculosis, pneumonia, other infections and parasites tended to be higher for American Indians in the latter years of the study

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TABLE 1—Infection-related Mortality among New Mexico Males, by Ethnic Group, 1958–87 (age-adjusted rates per 100,000)

	1958–62	1963–67	1968–72	1973–77	1978–82	1983–1987	% change in rates from 1958 to 1987
Tuberculosis							
Hispanic	13.4	9.6	5.1	3.8	3.0	1.1	−91.8
American Indian	40.5	40.5	27.3	23.5	10.7	11.7	−71.1
Other White	9.4	6.5	3.6	2.0	1.0	1.1	−88.3
US White*	7.8	5.3	3.3	1.2	1.2	0.9	−88.5
US Black*	22.6	16.9	12.1	7.1	6.1	4.9	−78.3
Other Infections and Parasites							
Hispanics	17.6	13.8	9.6	6.7	5.8	10.7	−39.2
American Indian	39.7	44.6	28.0	20.8	15.4	18.3	−53.9
Other White	5.7	4.7	4.9	4.5	6.0	8.4	+47.4**
US White	5.5	4.9	5.6	5.6	7.9	12.4	-125.5**
US Black	16.4	13.2	13.9	14.3	18.8	28.9	+76.2**
Influenza							
Hispanic	4.5	2.5	3.0	4.9	1.1	0.1	−97.8
American Indian	8.1	2.2	0.9	5.9	0.0	0.0	−100.0
Other White	2.9	2.3	3.9	4.4	0.9	0.5	−82.3
US White	4.6	1.2	2.1	2.3	1.1	0.7	−84.8
US Black	11.2	1.9	3.3	1.7	0.4	0.5	−95.5
Pneumonia							
Hispanic	50.7	44.5	56.9	52.2	36.3	33.9	−33.1
American Indian	88.5	64.5	81.2	89.4	48.3	51.6	−41.7
Other White	34.8	34.3	41.0	42.3	24.8	34.0	−2.3
US White	38.3	36.6	35.1	30.1	28.0	31.0	−19.1
US Black	72.6	63.8	61.6	44.3	40.6	39.9	−45.4
Kidney Infection							
Hispanic	4.5	5.2	4.6	1.6	1.0	0.1	−97.8
American Indian	10.6	8.6	4.5	1.0	2.5	1.5	−85.8
Other White	3.5	4.0	3.0	1.9	0.7	0.6	−82.3
US White	4.6	5.3	3.8	1.9	0.9	0.6	−87.0
US Black	11.0	12.6	9.1	3.8	1.7	1.1	−90.0
Meningitis							
Hispanic	1.2	1.2	0.8	2.1	1.0	0.4	−66.7
American Indian	7.0	10.8	8.2	1.6	2.5	0.5	−92.9
Other White	1.4	1.6	0.5	1.4	0.6	0.3	−78.6
US White	0.7	0.7	0.8	0.8	0.7	0.5	−28.6
US Black	2.3	3.2	2.3	2.2	1.8	1.3	−43.5

*Age-adjusted rates calculated at the midpoints of the time intervals.

**Rate increases primarily reflect increases in coding of deaths from zoonotic and other bacterial diseases, septicemia, and mycoses.

than for other Whites in the period 1958–1962, 30 years earlier.

Because of elevated mortality rates for infectious causes among young children in New Mexico, we examined age-specific mortality for pneumonia, meningitis, other infections, and parasites by sex and ethnic group among children ages 0–4 years. These data show trends similar to the age-adjusted data for Hispanics, American Indians, and other Whites in the state, with the highest mortality rates for the American Indian children (Table 3).

Discussion

Our data show that New Mexico's three major ethnic populations followed the national trend[19] of decreasing infectious disease-related mortality, but also indicate substantial ethnic differences in these mortality rates. These differences in mortality rates may be influenced by the disparate cultural beliefs and medical practices in these ethnic groups. Many Navajo, Pueblo, Zuni, and Apache people prefer traditional medicine men to modern medical treatment,[3,4] and some Hispanics use traditional curanderos.[20,21] Any consequent delay in seeking modern medical attention may increase mortality rates. Lack of health care availability can also affect mortality rates. In New Mexico, approximately one-third of Hispanics are medically uninsured, and many of the predominantly Hispanic rural counties are medically underserved.[22] Although the Indian Health Service provides care at no cost to the American Indians, long distances and unimproved roads also lessen health care availability, and have been demonstrated to be a risk factor for Navajo children presenting with infectious diseases.[23]

Our analyses of vital data have certain limitations; these potential sources of bias have been discussed in detail in prior publications.[9,10] In addition, these analyses cannot determine whether the incidence of infections is higher in the minority than in the White majority populations, or whether the incidence of infections is comparable, but the case-fatality ratios are higher in minority populations. Few published studies have directly addressed these issues from a population-based perspective on infectious disease incidence and mortality in the United States. The data from this study emphasize the importance of targeting the state's high-risk minority populations in provision of timely and appropriate medical care to reduce the high rates of mortality from infectious causes in New Mexico.

TABLE 2—Infection-related Mortality among New Mexico Females, by Ethnic Group, 1958–87 (age-adjusted rates per 100,000)

	1958–62	1963–67	1968–72	1973–77	1978–82	1983–1987	% change in rates from 1958 to 1987
Tuberculosis							
Hispanic	9.8	4.1	2.2	2.4	1.5	1.0	−89.8
American Indian	38.0	30.1	15.8	12.8	11.6	10.2	−73.2
Other White	3.7	1.7	0.5	0.7	0.8	0.6	−83.8
US White*	2.4	1.5	1.0	0.7	0.4	0.4	−83.3
US Black*	8.8	6.1	4.4	2.7	2.2	1.7	−80.7
Other Infections and Parasites							
Hispanics	11.9	7.4	7.6	6.4	6.7	6.9	−42.0
American Indian	37.0	38.3	22.5	20.9	12.2	4.5	−87.8
Other White	5.7	4.7	4.9	4.5	6.0	5.7	0
US White	3.5	3.2	4.3	3.9	5.6	8.1	+130.4**
US Black	9.3	8.5	10.0	10.7	13.6	19.1	+105.4**
Influenza							
Hispanic	2.8	3.1	4.7	4.6	0.6	0.2	−92.9
American Indian	4.9	1.5	2.5	3.1	0.7	0.8	−85.4
Other White	2.8	1.2	1.8	3.6	0.7	0.6	−78.6
US White	3.5	1.0	1.5	1.9	1.1	0.8	−77.1
US Black	9.0	2.1	2.4	1.2	0.4	0.3	−96.7
Pneumonia							
Hispanic	38.2	32.5	39.9	33.8	22.4	21.9	−42.7
American Indian	65.8	55.0	61.8	47.2	36.9	28.0	−57.4
Other White	19.2	18.7	25.4	26.0	17.0	21.3	+10.9
US White	21.3	22.4	20.8	16.9	17.3	19.8	−7.0
US Black	43.7	44.0	40.2	20.9	19.2	19.9	−54.5
Kidney Infection							
Hispanic	2.9	6.0	3.8	1.0	0.4	0.7	−75.9
American Indian	12.6	15.4	7.7	4.0	0.0	2.1	−83.3
Other White	2.6	3.6	3.7	1.6	0.7	0.6	−76.9
US White	3.7	4.6	3.5	1.9	1.1	0.7	−81.1
US Black	10.1	10.9	7.9	3.3	1.7	1.0	−90.1
Meningitis							
Hispanic	1.1	1.0	1.3	0.7	0.8	0.3	−72.7
American Indian	7.6	8.7	4.4	5.4	3.5	0.3	−96.1
Other White	0.5	1.1	1.0	0.8	0.3	1.2	+140.0
US White	1.0	1.0	0.6	0.6	0.5	0.4	−60.0
US Black	3.1	2.4	1.8	1.5	1.3	1.0	−67.7

*Age-adjusted rates calculated at the midpoints of the time intervals.

**Rate increases primarily reflect increases in coding of deaths from zoonotic and other bacterial diseases, septicemia, and mycoses.

TABLE 3—Infection-related Mortality among Children Ages 0–4 Years in New Mexico, 1958–87, by Cause and Ethnic Group (age-specific rates per 100,000)

	1958–62	1963–67	1968–72	1973–77	1978–82	1983–1987	% change in rates from 1958 to 1987
Pneumonia							
Males							
Hispanic	131.3	71.0	52.8	23.9	10.5	5.4	−95.9
American Indian	299.8	226.1	150.7	50.0	15.4	25.5	−91.5
Other White	46.0	36.8	30.7	22.8	6.7	4.0	−91.3
Females							
Hispanic	104.2	66.4	46.1	20.4	4.3	7.1	−93.2
American Indian	297.2	151.4	96.4	32.9	10.8	5.7	−98.1
Other White	48.1	35.2	25.6	21.0	6.6	6.1	−87.3
Meningitis							
Males							
Hispanic	5.8	6.4	3.6	10.0	6.5	0.8	−86.2
American Indian	67.3	77.7	59.9	6.7	12.9	5.5	−91.8
Other White	10.7	16.2	3.0	7.3	4.3	0.8	−92.5
Females							
Hispanic	5.1	8.5	9.4	1.7	4.1	2.3	−54.5
American Indian	50.5	65.8	29.6	29.6	9.2	0.0	−100.0
Other White	3.2	7.4	3.2	5.7	2.8	0.0	−100.0
Other Infections and Parasites							
Males							
Hispanic	110.8	60.7	26.8	16.5	16.0	7.6	−93.1
American Indian	323.8	290.6	166.8	80.0	30.5	21.7	−93.3
Other White	22.1	18.5	13.2	11.0	4.6	2.6	−88.2
Females							
Hispanic	98.5	39.9	30.7	18.7	11.6	2.6	−97.4
American Indian	246.9	197.0	98.0	72.4	16.1	3.1	−98.7
Other White	23.1	11.2	12.7	11.5	11.0	5.7	−75.3

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Feasibility of a Telephone Survey to Study a Minority Community: Hispanics in San Francisco

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Abstract: In two random digit dialing surveys conducted among Hispanics using a modified Mitofsky-Waksberg procedure, we found low refusal rates (4.7% and 3.1%), low assumed noncontact rate (14.0% and 18.3%), and high response rates (88.6% and 88.4%) with limited investment in time (1.58 hours and 1.66 hours per completed interview). These results suggest that Hispanics are willing to participate in telephone surveys and that this method may be feasible and useful for research and evaluation purposes. (*Am J Public Health* 1990; 80:323–326.)

Introduction

Surveys conducted over the telephone can be expected to be particularly useful to study minorities since they provide researchers with an efficient method to collect data with acceptable response rates and low interviewer bias.[1–4] Telephone surveys also allow the use of moderately long interview schedules regardless of respondents' reading and writing abilities. In addition, telephone surveys generally cost less than personal interviews. On the other hand, several investigators have suggested there are problems in identifying and sampling Hispanics; obtaining respondent cooperation; designing valid interview schedules; and controlling biases that may be introduced due to the ethnicity of the interviewer and the language used in the interview.[5–7]

The absence of a telephone at home among a significant proportion of potential respondents can be expected to decrease the representativeness of a sample and limit the generalizability of the results of a survey. Most individuals in the United States have a telephone in their household.[4] In San Francisco, for example, the 1980 Census found that 92 percent of self-identified Hispanic households in those census tracts with 400 or more Hispanics had at least one telephone compared with 94 percent for non-Hispanics. In other major cities telephone ownership among Hispanics averages 85 percent.[8]

We used the random digit dialing procedure for telephone surveys suggested by Mitofsky and Waksberg[9] which eliminates the use of nonworking banks (prefixes) and avoids the use of directories (eliminating the problem of excluding

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unlisted numbers) while guaranteeing equal probability of selection for all residential telephone numbers.[1] A recent study with Blacks[10] found this approach particularly efficient in obtaining samples from highly clustered subgroups as is the case of many ethnic/racial minorities living in urban areas of the United States. This paper describes two surveys in which the Mitofsky-Waksberg approach was used to sample Hispanics.

Methods

The surveys were designed to obtain estimates of the prevalence of cigarette smoking[11] and to serve as baselines for the evaluation of a community intervention designed to lower cigarette smoking prevalence among Hispanics in San Francisco. Each interview lasted approximately 15 minutes and included approximately 60 questions dealing with smoking behavior, standard socio-demographic variables, and an acculturation scale.[12] The interviews were conducted in the language of choice of the respondents; 67 percent and 69 percent preferred to use Spanish in the first and second studies, respectively. Interviewers were bilingual (English and Spanish) men and women who received at least 12 hours of training, including role playing and supervised interviewing, before they participated in the study. Training of interviewers was based on procedures for telephone interviewing developed by the Minnesota Heart Health Program and on materials from the Survey Research Center of the University of Michigan.[13]

A household was judged to be eligible when the respondent identified himself/herself as Hispanic, or identified the majority of the residents in the household as "Hispanics." Hispanic ethnicity was ascertained by the respondents' answer to the question "Do you or most of the people in your household consider yourselves to be Hispanics, Blacks, Asians, or Caucasians?" The response to this question was later checked with the respondents' and their parents' reported place of birth and with the respondents' identification with one of the Hispanic subgroups.

Prefix Selection

We selected prefixes that covered census tracts with the largest concentration of Hispanics, according to 1980 census data. This identified 27 tracts with 10 percent or more Hispanic residents. These tracts represented 66.6 percent (N = 55,541) of all Hispanics (N = 83,373) in the city of San Francisco.

In each census tract with high concentrations of Hispanics, three city streets were randomly selected to locate telephone prefixes assigned to residences. This process produced 49 three-digit prefixes out of the 121 prefixes in use in San Francisco. These procedures were conducted separately for each study and in both cases the same 49 three-digit prefixes were identified and used.

Stage One Sampling

A random list of 2,000 different four-digit numbers was generated by a computer and the numbers were randomly assigned to the 49 prefixes. These seven-digit telephone numbers were then assigned to interviewers. When a Hispanic household was reached, the first five digits of that telephone number became a Primary Sampling Unit (PSU). During Study I, a total of 4,326 working numbers were called as part of this first stage in the sampling procedure yielding 259 PSUs. Study II produced 355 PSUs out of the 4,680 working numbers called. Tables 1 and 2 present a summary of the outcome of the calls made in this stage during both studies.

TABLE 1—Summary of Telephone Calls for Study I

Outcome	Stage I N (%)	Stage II N (%)	Total N (%)
Total working numbers called	4,326 (100)	12,107 (100)	16,433 (100)
Eligible			
Completed	286 (6.6)	1,383 (11.4)	1,669 (10.2)
Refusals	20 (.5)	63 (.5)	83 (.5)
Ineligible			
Business	852 (19.7)	1,704 (14.1)	2,556 (15.6)
Ethnicity	2,232 (51.6)	8,605 (71.1)	10,837 (65.9)
Age	35 (1.0)	114 (1.0)	149 (1.0)
Lost to survey	901 (20.1)	238 (2.0)	1,139 (6.9)
Total screened	3,425 (79.9)	11,869 (98.0)	15,294 (93.1)
Response rate	(74.1)	(93.8)	88.6
Refusal rate	(6.5)	(4.3)	(4.7)
Assumed noncontact rate	(42.4)	(4.0)	(14.0)

TABLE 2—Summary of Telephone Calls for Study II

Outcome	Stage I N (%)	Stage II N (%)	Total N (%)
Total working numbers called	4,680 (100)	16,823 (100)	21,503 (100)
Eligible			
Completed	365 (7.8)	1,694 (10.1)	2,059 (9.6)
Refusals	11 (.2)	55 (.3)	66 (.3)
Ineligible			
Business	1,245 (26.6)	3,347 (19.9)	4,592 (21.4)
Ethnicity	2,735 (58.4)	9,942 (59.1)	12,677 (59.0)
Age	40 (.9)	167 (1.0)	207 (1.0)
Lost to survey	284 (6.1)	1,618 (9.6)	1,902 (8.8)
Total screened	4,396 (93.9)	15,205 (90.4)	19,601 (91.2)
Response rate	(91.3)	(87.5)	(88.4)
Refusal rate	(2.9)	(3.1)	(3.1)
Assumed noncontact rate	(15.9)	(18.8)	(18.3)

Stage Two Sampling

A computer-generated list of 99 random two-digit numbers was used to produce the suffixes for each of the PSUs. Each interviewer continued using a given PSU until eight Hispanic households were reached including completed interviews or refusals, or until the list was exhausted including four follow-up calls made within a four-week period to every unanswered number. Tables 1 and 2 present a summary of the outcome of the numbers called in this stage during both studies. A total of 92 PSUs (36 percent) were exhausted before reaching eight Hispanic households during the first study. The corresponding figure for the second study was 95 (25 percent).

Interviewee Eligibility

Within a Hispanic household, the respondent was selected by asking for the resident who had most recently celebrated a birthday and who was between 18 and 65 years of age. Approximately 8 percent of otherwise eligible households failed to meet this criterion (Tables 1 and 2).

A total of 1,669 respondents were interviewed during Study I (Fall 1986). Of these, 44 percent were males and the majority of the respondents (69 percent) were foreign-born with 60 percent of them reporting being born in Central America, 29 percent in Mexico, and the remainder in South America, Cuba, or Puerto Rico. The respondents' mean age was 35.5 years and they reported a mean number of years of education of 11.6. In Study II (Summer 1987), 2,059 Hispanics were interviewed (42 percent men). The respondents' mean age was 35.5 years while the mean number of years of education was 11.6. The majority (70 percent) were born outside the United States, 60 percent of them in Central America, 28 percent in Mexico, and the remainder in either Cuba, Puerto Rico, or South America.

Results

Response/Refusal Rates

Tables 1 and 2 show the response and refusal rates for both studies. We have adapted Kviz's[14] definition of response rate which is the proportion of completed interviews in terms of the number of eligible respondents (defined here as Hispanics between 18 and 65 years of age). This calculation assumes that those not contacted after four follow-up calls ("lost to survey") were eligibles.[15] This proportion was assumed to be the same as the proportion of eligibles found among telephone numbers that were fully screened. The Stage I response rate, assumed that the proportion of eligibles found among the screened calls 306/3425 (8.9 percent) would be similar to that among the 901 numbers that were "lost to survey." Adding the 80 assumed eligible numbers yielded a response rate of 286/386 (74.1 percent). Response rates for the other stages of both studies were calculated in the same manner noted here and show an overall response rate of 88.6 percent for Study I and 88.4 percent for Study II.

The refusal rates were defined[1] as the proportion of eligible respondents (Hispanics between 18 and 65 years of age) who were contacted but declined to be interviewed. Refusal rates were only 4.7 percent for Study I and 3.1 percent for Study II.

The Assumed Noncontact Rate[1] is an index of the accessibility of respondents or the probability of researchers locating respondents. This index is made up of the proportion of respondents not contacted after four follow-up calls ("lost to survey" in both tables) in relation to the total number of possible eligible respondents. Given that a significant proportion of telephone numbers in San Francisco are unlisted (38 percent according to Pacific Bell), it is not appropriate to estimate the true number of eligible households from an alphabetic telephone directory. Thus, in computing the assumed noncontact rate we are assuming that 25 percent of those respondents "lost to survey" were eligible Hispanics. This assumption reflects twice the number of Hispanics reported to be living in San Francisco in 1985 and therefore produces a very conservative computation of nonresponse rates. In Study I for example, 25 percent of the 1,139 "lost to survey" (N = 285) represent 14 percent of the total number of assumed Hispanic numbers: completed interviews (N = 1,669) plus refusals (N = 83) plus lost to survey (N = 285). The comparable noncontact rate for Study II is 18.3 percent. Figures taken from Wiseman and McDonald's[16] review of the literature show a median noncontact rate of 39.1 percent.

Cost of Survey

Based upon the time spent by all interviewers, we estimated costs of obtaining a completed interview. On the average, interviewers needed to spend 1.58 hours in order to obtain a completed interview during Study I and 1.66 hours during Study II. Each of the completed interviews cost approximately $15 based on paid interviewer and supervisor salaries but excluding overhead, training time, telephone lines, investigators' time, and data management. Interviewers were paid $7.90 per hour in this study.

Discussion

These results suggest that it is feasible and efficient to use random digit dialing techniques to obtain a sample of Hispanics. The similarity in the outcomes of the two studies imply that the findings are fairly stable across time.

The high completion rates and the correspondingly low refusal rates obtained in these surveys compare favorably with those found in studies of other Whites, which range between 66 percent in a study in Chicago[4] to 73.5 percent in a regional survey.[17] The low refusal rates (4.7 percent in Study I and 3.1 percent in Study II) are lower than those of other studies. Wiseman and McDonald[16] in their review of 182 studies using telephone surveys found a median refusal rate of 28 percent while Frey[1] reports refusal rates ranging between 20–25 percent. Low refusal rates among Hispanics have also been reported in a study conducted in Texas[18] where a telephone survey of a sample drawn from Spanish-surnamed individuals produced a refusal rate of 8.8 percent.

The high rate of participation in these surveys by Hispanics was not due to advance publicity about the study. Respondents were informed of the purpose of the study only at the time of the interview. One possible explanation for these results is the presence of a cultural characteristic that promotes cooperation and that dictates positive interpersonal relations—the social script of "simpatia."[19] This cultural norm would move Hispanic respondents to agree to be interviewed in order to maintain a fluid and positive social relationship with the interviewer when the relationship is perceived as nonthreatening and as sanctioned by a respected local institution (the study was introduced as being conducted by a state university in San Francisco).

The generalizability of our results may be limited by the following specific concerns:

- The sampling frame used excluded approximately one-third of the San Francisco Hispanic population according to the 1980 census. Hispanics who are living in census tracts excluded in this study (less than 10 percent Hispanic) may behave differently when asked for an interview over the telephone.
- At least 8 percent of San Francisco Hispanic households may not have a telephone at home and these persons are also excluded from the sample.
- The approach utilized in choosing the prefixes may not produce all of the relevant ones although a third study survey recently concluded by the authors has again identified the same 49 prefixes as the first two studies.
- Finally, San Francisco Hispanics (Central American and Mexican in this survey) may differ from Hispanics residing elsewhere.

Future studies should compare samples gathered with this method with census data in order to address the representativeness of a sample of minority-group respondents obtained through a random digit dialing technique such as the Mitofsky-Waksberg approach. Unfortunately, this analysis cannot be done at the present time due to the high undercount of Hispanics in the 1980 census and the later significant migrations from Cuba and Central America.

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Sex-specific and Race-specific Hip Fracture Rates

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Abstract: Sex-, race- and age-specific hip fracture rates were determined using Health Care Financing Administration data for Medicare-reimbursed hip fracture hospitalizations from 1980 to 1982. Rates were highest in White women, lowest in Black men, and intermediate in White men and Black women. Proportions of hip fracture patients dying during hospitalization and those discharged to nursing homes, respectively, were: White men (10.5%; 49%); Black men (9.3%; 32%); White women (5.0%; 54%); and Black women (8.2%; 30%). (*Am J Public Health* 1990; 80:326–328.)

Introduction

In 1986 in the United States, there were in excess of 250,000 hospital admissions for hip fracture.[1] Elderly individuals who sustain hip fractures experience increased morbidity[2–7] and mortality.[8–12] In an earlier paper, Farmer, *et al*,[13] reported that age-specific hip fracture rates were higher in White women than in White men, Black women, and Black men, but the differences among age-specific rates for White men, Black men, and Black women could not be estimated precisely.

In this paper we report age-, sex-, and race-specific hip fracture rates extended through age 85 years. In addition, we present findings for proportions of patients dying during hospitalization, and discharged to nursing homes, not previously reported for Blacks.

Methods

Numbers of Medicare-reimbursed hospitalizations for hip fracture in Illinois residents over 65 years of age were obtained from data collected by the Professional Standards Review Organization in Illinois during 1980, 1981, and 1982. The Uniform Hospital Discharge Format was used in the data collection, and the data tape was made available through the Health Care Financing Administration (HCFA). Hip fracture was defined according to the ninth revised International Classification of Diseases[14] categories 820.0–820.9: transcervical fractures, closed and open; pertrochanteric fractures, closed and open; and fractures of unspecified parts of the neck of the femur, closed and open. Populations at risk for hip fracture were obtained from the 1980 United States Census figures for the Illinois population by age, race and sex.

Mean numbers of hip fractures for 1980, 1981, and 1982 were used to calculate the sex-, race- and age-specific annual incidence rates for hip fracture. Miettinen test-based confidence intervals were calculated for the hip fracture relative risks.[15] Age-adjusted odds ratios for death and nursing home placement were determined using the Statistical Analysis System multiple logistic regression program.[16] Race- and sex-specific means for days of hospital stay were adjusted for age with the Statistical Analysis System linear regression program.[17]

Results

Age-group specific hip fracture rates were highest for White women, and lowest for Black men in all age groups after age 70 years (Figure 1). The rates in Black women are

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similar to those in White men, and intermediate to those observed for White women and Black men. Hip fracture rates began to increase exponentially after age 70 years in White women, after age 75 in White men and Black women, and at approximately 85 years of age in Black men. Hip fracture rates in White women were approximately twice those in White men from age 65 to 85 years.*

In-hospital hip fracture death rates were highest for White men (10.5%), followed by Black men (9.3%), Black women (8.2%), and White women (5.0%). Age-adjusted odds ratios for death among hip fracture cases are shown in Table 1. After age-adjustment, in-hospital death rates were twice as high for White men as for White women. Black men had only slightly higher death rates than Black women.

The means for lengths of hospital stays were: Black men (28.0 days), Black women (28.2 days), White men (24.2 days), and White women (23.1 days). Discharge status of survivors is shown in Table 2. Both Black men and women were discharged to nursing homes less frequently than White men and women. After age-adjustment, the odds ratio for nursing home placement in White women compared with Black women was 2.70 (95% CI: 2.21, 3.30), and in White men compared with Black men 1.93 (95% CI: 1.16, 2.61).

Discussion

These data are consistent with earlier observations showing highest age-specific hip fracture rates for White women.[13] However, our findings also indicate higher age-specific rates in women than in men among both Blacks and Whites, although the sex difference occurred at older ages in Blacks than in Whites, as did the race-related differences. These findings are consistent with the age-, sex- and race-specific hip fracture rates recently reported for hip fracture hospitalizations in California.[18]

Black men and women may have lower hip fracture rates because, as adults, they exhibit higher levels of bone mass attained during growth and development.[19–25] In addition, race-related differences in the Vitamin D endocrine system suggest that Blacks are protected from the bone-resorptive effects of parathyroid hormone,[26] and reports of lower bone turnover rates in Blacks may result in slower rates of age-related bone mass than occurs in Whites.[27] The race difference in women is compatible with evidence that obesity is more common in Black women which may increase their extra-ovarian production of estrogen,[28] and that Black women may experience lower rates of menopause-associated bone loss than White women.[29]

While risk for falling, in addition to bone mass, influences hip fracture incidence,[30,31] the extent to which falling risk contributes to race- and sex-differences in hip fracture rates remains unclear. Elderly men are reported to fall less frequently than elderly women.[32] While Black women are reported to have greater muscle mass than White women,[33–35] it is not known if muscle mass acts independently of its documented influence on bone mass, to influence falling risk.

The finding of higher mortality rates in White men than in women is consistent with observations from earlier studies.[8–12] The most comparable race-specific mortality data were reported for small numbers of non-White hip fracture patients in which the risk for mortality at three months post hip fracture in non-Whites versus Whites was 1.5 (95% CI: 0.6, 3.6).[12] In addition to length of hospital stay, in-hospital mortality rates are also influenced by the numbers and severity of co-existing illnesses.[12] In the absence of comorbidity and severity of illness data in this analysis, interpretation of the in-hospital mortality data reported is difficult.

Hospital lengths of stay observed in this study are for the period prior to the implementation of the diagnosis-related group (DRG)-based prospective payment system (PPS), and are consistent with those reported by Fitzgerald, *et al*,[36] which also noted that the shorter lengths of stay after the PPS (12.6 days) were associated with fewer in-hospital physical

TABLE 1—Age-adjusted Odds Ratios for Death during Medicare-reimbursed Hospitalization for Hip Fracture, Illinois, 1980–82

	Age-adjusted Odds Ratios for Death	95% CI*
White men vs White women	2.41	2.17, 3.00
Black men vs Black women	1.18	0.53, 2.35
White men vs Black men	1.02	0.63, 2.23
Black women vs White women	1.64	1.09, 2.47

*Miettinen, test-based confidence intervals.[15]

TABLE 2—Discharge Status of Survivors and Duration of Stay for Medicare-reimbursed Hospitalization for Hip Fracture, Illinois, 1980–82.

	White		Black	
	Men (N = 3575)	Women (N = 14,768)	Men (N = 204)	Women (N = 523)
Nursing home (%)	49	54	32	30
Skilled (%)	73	73	69	66
Intermediate (%)	23	24	29	30
Other (%)	4	3	2	4
Home (%)	46	43	65	66
Self-care (%)	91	90	94	93
Home-health (%)	9	10	6	7
Transferred to another hospital (%)	4	3	2	3
Left hospital against medical advice	[illegible]	[illegible]	1	1

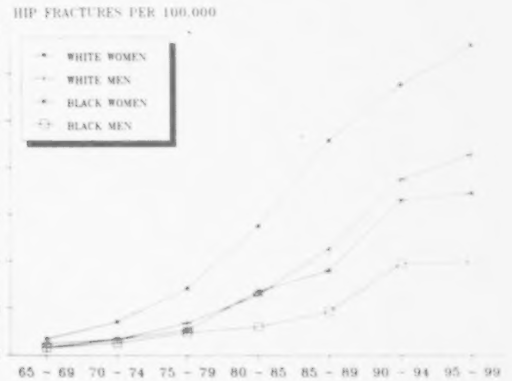

FIGURE 1—Comparison of Sex-specific and Race-specific Hip Fracture Incidence Rates by Age for Illinois Residents, 1980–82

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*Exact data available upon request from author.

therapy sessions, poorer ability to ambulate on hospital discharge, and higher rates of discharge to nursing homes.

The finding of lower proportions of Black men and women discharged to nursing homes is consistent with race-specific national data for nursing home use.[37]

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The Epidemiology of Vulvovaginal Candidiasis: Risk Factors

BETSY FOXMAN, PHD

Abstract: The effects of personal hygiene, sexual history, diet, and stress on the risk of vulvovaginal candidiasis were estimated from a case-control study of students attending a public university during 1986–87. Data from medical records and self-administered questionnaires were used to compare 85 cases to 1,245 other students using the Health Service, and to 113 subjects chosen from the total student population. Frequent sexual intercourse was the strongest risk factor (seven or more times a week versus none): OR = 4.3; 95% CI: 1.4, 12.9 (for cases versus Health Service controls). (*Am J Public Health* 1990; 80:329–331.)

Introduction

Vulvovaginal candidiasis, a common, irritating and frequently recurring infection, has been associated with broad spectrum antibiotic treatment, diabetes, and pregnancy.[1] Although factors such as diet high in sugar, using tampons or perfumed toilet paper, or wearing tight clothing are suspected to increase risk of vulvovaginal candidiasis,[1,2] few studies have examined these factors in a systematic fashion. In this study, I explored the relationships between personal habits, behaviors, and vulvovaginal candidiasis by analyzing a subset of data from a Health Service-based case-control study of urinary tract infection.[3]

Methods

Cases and controls were collected as part of a previously described study of urinary tract infection.[3] For that study, participants were recruited from women presenting with and without urinary symptoms to the University Health Service between September 1986 and December 1987. Prior to seeing a clinician, participants were given a self-administered questionnaire including questions about medical history, stress, clothing, diet, sexual activity, and birth control method during the past four weeks. The medical records of all participants recruited at the Health Service were reviewed to establish diagnosis and to confirm information on birth control method and history of other infections.

There were two control groups: women using the Health Service during the same time period as the cases (Health Service controls); and a sample stratified by year in school of 200 women enrolled at the University during January, 1988 (student population controls). Of these 200, 170 had complete and accurate addresses and 115 consented to participate. Student population controls completed essentially the same questionnaire distributed at the Health Service, but medical records were not reviewed.

Case and Control Definition

A case was defined as any woman who was recruited during a visit for which the diagnosis listed in her medical record was vulvovaginal candidiasis (n = 85). Health Service controls were women without a clinical diagnosis of vulvovaginal candidiasis recruited from the University Health Service. Population-based controls were limited to women who did not report current infection with either vulvovaginal candidiasis or a urinary tract infection. Both cases and controls were limited to non-pregnant women without urinary tract infection enrolled at the University.

Analysis

Exploratory analyses were conducted using multilevel contingency tables with both stratification and multivariate techniques to control for confounding.[4,5] Odds ratios (OR) and associated 95 percent confidence intervals (95% CI) were calculated from the logistic regression models as described by Schlesselman.[4] A separate model was fit for each factor of interest with age, frequency of sexual intercourse, number of previous infections with vulvovaginal candidiasis, and, for Health Service controls, month of enrollment, included as covariates. Related variables, such as birth control methods, were fit in a single model. All factors of interest were considered as potential confounders of other factors but only those variables that made a difference were included as covariates in the final models. For example, the OR for feminine hygiene use for cases compared to Health Service controls was calculated from a logistic regression model with presence of vulvovaginal candidiasis as the dependent variable, and age, frequency of sexual intercourse, number of prior infections with *Candida*, month of enrollment, and feminine hygiene use as independent variables.

Results

A total of 85 women met the study case definition; 1,245 Health Service and 113 population-sampled participants met the control definition. Study participants were similar to women students enrolled in the University with respect to year in school, age, ethnic group, and marital status. Cases and controls ranged in age from 17 to 72 (mean = 21.8 years). Age, year in school, and marital status were similar for cases and controls. Self-reported history of previous vulvovaginal candidiasis was common among both cases (77.7 percent) and controls (Health Service: 41.2 percent; student population: 19.6 percent). Most previous infections occurred within the past 12 months (cases: 74 percent, Health Service: 60 percent; student population: 62 percent).

When controlled for the confounding effects of age, frequency of sexual intercourse, and month of enrollment, women reporting one previous infection, compared to none, were about twice as likely to have a current infection (Health Service: OR = 1.55; 95% CI: 1.35, 1.78; population: OR = 2.06; 95% CI: 1.58, 2.69), and those with two infections about three times as likely (Health Service: OR = 2.39; 95% CI: 1.81, 3.16; population: OR = 4.25; 95% CI: 2.49, 7.25).

Cases were twice as likely as controls to have been sexually active during the past four weeks; the odds ratio increased with increasing frequency of sexual intercourse for both control groups (Table 1). However, having a new partner or engaging in sexual intercourse with multiple partners during the past four weeks showed no association

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TABLE 1—Associations between Vulvovaginal Candidiasis and Selected Behaviors during the Past Four Weeks

Variables	Cases (n = 85)		Health Service Controls (n = 1245)			Population Controls (n = 113)		
	#	%	#	%	OR (95% CI)	#	%	OR (95% CI)
Weekly frequency of sexual intercourse during the past 4 weeks								
never	19	22.4	557	44.7	reference	68	60.2	reference
< 1 time	14	16.5	192	15.4	1.31 (1.10, 1.57)	15	13.3	1.44 (1.10, 1.90)
1–2 times	20	23.5	231	18.6	1.73 (1.22, 2.45)	11	9.7	2.08 (1.21, 3.60)
3–6 times	25	29.4	218	17.5	2.27 (1.34, 3.84)	17	15.0	3.00 (1.32, 6.82)
7+ times	7	8.2	47	3.8	2.98 (1.48, 6.00)	2	1.8	4.34 (1.45, 12.93)
Birth control method† (sexually active women only)								
No method	15	22.7	16	1.3	reference	9	20.9	reference
Oral contraceptives	35	53.0	373	55.4	1.11 (0.60, 2.02)	23	53.5	1.52 (0.61, 3.74)
Diaphragm	10	15.2	74	37.9	1.23 (0.40, 3.80)	6	14.0	1.67 (0.30, 9.44)
Spermicides	18	27.3	79	11.7	1.01 (0.43, 2.35)	9	20.9	1.40 (0.36, 5.53)
Other methods	21	32.8	264	40.0	0.95 (0.51, 1.78)	22	51.2	1.01 (0.38, 2.70)

NOTE: Odds ratios and their 95 percent confidence intervals (OR, 95% CI) were adjusted for frequency of sexual intercourse, previous infections with vulvovaginal candidiasis, age, and for Health Service controls, month of enrollment, using a logistic regression model (see methods).

†Categories not mutually exclusive. Percents do not sum to 100, as some women used more than one method during the previous four weeks.

with *Candida*. Moreover, no one birth control method was associated with infection (Table 1).

Student population controls were more likely to use colored toilet paper and take tub baths than cases, but Health Service controls reported similar exposures to cases (Table 2). Other suspected risk factors, direction wiped following toileting, menstrual protection, use of feminine hygiene products and clothing showed little association with vulvovaginal candidiasis (Table 2). Diet and stress also showed no association.

Discussion

To my knowledge, no previous controlled study has reported a relationship between sexual activity and vulvovaginal candidiasis although case reports suggest a correlation.[1] Sexual intercourse might facilitate movement of *Candida* into the vagina and result in local minor trauma, creating conditions suitable for yeast to invade tissues. Some authors have suggested that *Candida* can be sexually transmitted;[6] however, in this study neither beginning a new relationship nor having multiple sexual partners in the previous four weeks was associated with candidiasis after controlling for frequency of sexual intercourse.

Although similar with respect to demographic variables to all University students, it is possible that women using the Health Service are a select group. They were more likely to be sexually active and use contraceptives than the student population controls. The resulting selection bias might decrease the observed associations between sexual activity, or variables related to sexual activity, and candida when cases are compared to Health Service controls. Further, Health

TABLE 2—Associations between Vulvovaginal Candidiasis and Selected Behaviors during the Past Four Weeks

Variables	Cases (n = 85)		Health Service Controls (n = 1245)			Population Controls (n = 113)		
	#	%	#	%	OR (95% CI)	#	%	OR (95% CI)
Wipe back to front vs reverse	21	25.0	403	33.0	0.86 (0.51, 1.45)	59	52.2	0.34 (0.16, 0.71)
Infrequent bowel movements	13	15.3	148	11.9	1.26 (0.68, 2.31)	11	9.7	1.63 (0.60, 4.44)
Use colored toilet paper								
never	25	29.8	505	41.0	reference	62	54.9	reference
rare	30	35.7	372	30.2	1.07 (0.88, 1.31)	29	25.7	1.41 (0.71, 2.00)
some	17	20.2	215	17.5	1.14 (0.77, 1.70)	15	13.3	1.99 (0.99, 4.00)
frequently	11	13.1	118	9.6	1.22 (0.67, 2.33)	7	6.2	2.81 (0.99, 8.01)
always	1	1.2	22	1.8	1.31 (0.59, 2.91)	—	—	—
Menstrual protection								
napkins	11	13.1	214	17.2	reference	26	23.4	reference
tampons	31	36.9	384	30.8	1.19 (0.58, 2.43)	27	24.3	1.34 (0.48, 3.76)
tampons & napkins	39	46.4	578	46.4	1.20 (0.61, 2.38)	49	44.1	1.43 (0.56, 3.65)
Tub baths (ever)	33	39.3	433	35.3	0.90 (0.56, 1.45)	18	16.1	3.49 (1.54, 7.95)
Feminine hygiene products	15	17.7	76	6.1	1.81 (0.94, 3.49)	1	0.9	12.89 (1.39, 119.79)
Underwear fabric								
always cotton	23	27.1	335	26.9	reference	38	33.6	reference
mix of fabrics	57	67.1	820	65.9	0.94 (0.56, 1.57)	71	62.8	1.05 (0.50, 2.21)
synthetic only	5	5.9	90	7.2	0.99 (0.36, 2.75)	4	3.5	0.96 (0.16, 5.88)
Wearing tight clothing	6	7.1	105	8.4	0.70 (0.29, 1.69)	8	7.1	1.72 (0.48, 6.19)
UTI in past year	20	23.5	204	16.6	1.28 (0.72, 1.38)	7	6.2	2.27 (0.74, 7.03)

NOTE: Odds ratios and their 95 percent confidence intervals (OR, 95% CI), were adjusted for frequency of sexual intercourse, previous infections with vulvovaginal candidiasis, age, and for Health Service controls, month of enrollment, using a logistic regression model (see methods).

Service users may be more likely to have diseases associated with sexual activity and thus be more likely to adopt behaviors thought to limit infection. This may account for the differences in associations when cases are compared to student population rather than Health Service controls. Cases and controls completed the same questionnaires minimizing any misclassification with respect to exposure. Because case status was based solely on clinical diagnosis, undoubtedly some misclassification of disease occurred. However, the bias should have been at random, biasing the results toward no difference between groups.

This study showed a strong association between sexual intercourse and vulvovaginal candidiasis among non-pregnant women. Other suggested factors—wearing tight clothing, hose or synthetic underwear; type of menstrual protection; use of feminine hygiene sprays; and diet—were not associated with candidiasis. Most women with recurring vulvovaginal candidiasis tend to avoid all suggested risk factors,[1] which do not include sexual intercourse. Although these findings need confirmation, suggesting to women suffering recurring vulvovaginal candidiasis that they temporarily decrease frequency of sexual intercourse has more scientific basis than some of the other standard advice (wear cotton underwear), would do no harm, and might potentially avoid recurrence.

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Recurring Urinary Tract Infection: Incidence and Risk Factors

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Abstract: Urinary tract infection (UTI) is a common infection among young women, with a high recurrence rate. This study documents the six-month incidence of second UTI among a cohort of women with one initial UTI and the factors associated with recurrence. Among the cohort of 113 women, 30 (26.6 percent) experienced at least one culture-confirmed recurrence within the six months following initial infection. The presence of hematuria and urgency as symptoms of initial infection were the strongest predictors of second infection. Behavioral factors associated with initial infection (frequency of sexual intercourse, diaphragm use, and voiding after sexual intercourse) did not distinguish between women who would and would not experience a second UTI during the six-month follow-up period. (*Am J Public Health* 1990; 80:331–333.)

Introduction

Urinary tract infection occurs frequently among young women, with over one-fourth having a second infection.[1] Why otherwise healthy women with no known anatomical abnormalities have multiple recurrences is a subject of some controversy, although both host and agent characteristics have been suspected. However, most authors would agree with Kunin's statement that the risk of a second infection is greater than the first, and that the first "sets the stage" for recurring infection.[2]

Methods

All cases of first UTI among college women enrolled in a previously described case-control study* were eligible to participate in the follow-up study. A UTI was defined as the presence of significant bacteriuria (≥10,000 colonies of a single bacteria per milliliter urine) and one or more urinary symptoms: hematuria, dysuria, frequency, urgency, or suprapubic pressure. Urinalyses and urine cultures were performed on all women presenting with urinary symptoms using standard techniques.* History of UTI was assessed by review of medical records and self-report.

Prior to enrollment, participants completed a self-administered questionnaire which contained a consent form approved by the University of Michigan Institutional Review Board and questions on medical history, stress, clothing, diet, sexual activity, and birth control method. Responses to questions regarding birth control method, sexual activity, history of UTI, and demographics and the presence of UTI were confirmed by review of medical records. At the end of the six-months follow-up, participants' medical records were again reviewed to record the number and timing of UTIs treated.

Symptoms (frequency, dysuria, nocturia, urgency, hematuria, fever, chills, back or flank pain, and any other symptoms specified by the respondent), costs of treatment, measures of disease impact, and whether a woman had changed her health habits following the UTI were assessed via a self-administered questionnaire sent to all participants two weeks following enrollment.

Since bacteria normally found in the intestine cause the majority of UTIs in this age group, a woman may be

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*Foxman B, Chi JW: Health behavior and urinary tract infection 1990; J Clinic Epidemiol (in press).

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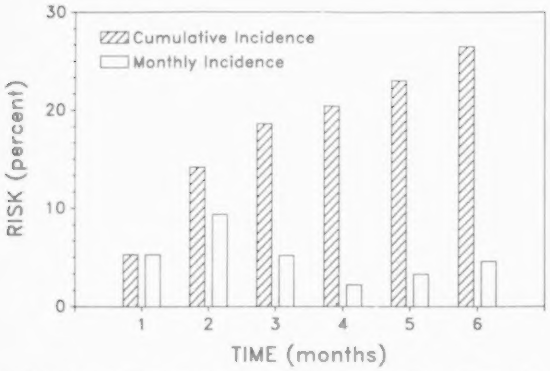

FIGURE 1—Monthly and Cumulative Incidence of Second Urinary Tract Infection (UTI) among a Cohort of 113 Women with One Prior UTI

re-infected with identical bacteria making it difficult to distinguish relapse from recurrence even in the presence of sophisticated bacteriologic techniques. As such techniques were not used in this study, I defined a recurrent urinary tract infection as the presence of ≥ 10,000 col/ml of a single bacteria and one or more urinary symptoms occurring at least two weeks after the initial infection. This may lead to some misclassification of outcome; however, as long as the classification is not associated with the exposure(s) of interest the resulting measures of association should be biased toward the null hypothesis of no difference.

The associations between hypothesized risk factors and repeat UTI were studied using multilevel contingency tables, with both stratification and multivariate techniques to control for confounding.[3] I calculated six-month cumulative incidences of UTI, risk ratios (RR), and Mantel-Haenszel summary RRs. Exact 95 percent confidence intervals (CI) were calculated around each RR using the method of Gart.[4]

A three-level symptoms scale, (0 = no hematuria or urgency, 1 = urgency alone, 2 = hematuria and urgency) was developed and used both as a dependent and independent variable in the analysis. When used as a dependent variable, I tested the hypothesis that the relationship was linear and the slope zero with a chi square test for trend.[5] Linear regression models with number of culture confirmed repeat UTIs and with the symptom scale as dependent variables were also fit to the data.

Results

The study cohort consisted of 113 women ages 17 to 39 (mean 21.3 years) with a first UTI. None of the participants had pyelonephritis, nor concurrent infection with gonorrhea, herpes, or chlamydia. The majority were White (83.8 percent), single (90.1 percent), and did not have a private physician (other than at the Health Service) within driving distance of their current residence (71.2 percent). Most had sexual intercourse during the four weeks prior to their first UTI (94.6 percent).

The six-month risk of a second UTI was 26.6 percent; the risk of a third UTI, 2.7 percent. Figure 1 shows the distribution of second infections; the median time to recurrence was approximately two months (56 days).

The presence of hematuria and/or urgency as symptoms of first infection was the strongest predictor of second infection (Table 1). Although the confidence intervals around the risk ratios include 1.0 (no difference), a significant relationship between symptoms and second infection was observed using a linear regression model including age as an independent variable (linear regression coefficient: 0.149, SE = 0.064). Other symptoms—fever, chills, flank or back pain, frequency, nocturia, or dysuria—were not associated with second infection.

Frequency of sexual intercourse, diaphragm use, and voiding around sexual intercourse were only weakly associated with second infection (Table 1). The trend in decreasing proportion of second infections with voiding after sexual intercourse is of interest (linear regression coefficient: −0.033; SE = 0.033).

Neither frequency of sexual intercourse, diaphragm use, nor voiding after sexual intercourse were associated with the presence of urgency and hematuria at first infection. However, increasing age was negatively associated with increas-

TABLE 1—Factors Associated with Second Urinary Tract Infection among College Women

Factors	Number	Percent Recurring	RR (95% CI)
*Symptoms of initial infection**			
Symptom scale			
no urgency or hematuria	22	13.6	1.0 (reference)
urgency alone	38	21.1	1.5 (0.5, 5.2)
urgency and hematuria	47	38.3	2.8 (0.9, 8.5)
Behaviors during four weeks prior to initial infection			
Frequency of sexual intercourse†			
<1 times per week	20	25.0	1.0 (reference)
1–2 times per week	38	21.0	0.8 (0.3, 2.2)
3 or more times per week	52	30.8	1.2 (0.5, 2.9)
Birth control method (sexually active women only)			
Oral contraceptives	47	25.5	1.0 (reference)
Diaphragm use	22	31.8	1.2 (0.6, 2.7)
Other	36	25.0	1.0 (0.5, 2.1)
Voiding after sexual intercourse (sexually active women only)			
Never/rare	36	33.3	1.0 (reference)
Sometimes	34	26.5	0.8 (0.4, 1.8)
Frequently/always	35	20.0	0.7 (0.3, 1.5)

*Two women had hematuria alone, four had missing data.
†Three women had missing data.

ing symptoms (linear regression coefficient: −0.012; SE = 0.013).

Discussion

In this study, factors observed to be associated with UTI in previous studies[6–9]—frequency of sexual intercourse, diaphragm use, and voiding after sexual intercourse—did not predict recurrence in the next six months. The strongest predictors were characteristics of the infection itself: the presence of hematuria and urgency. Indeed, the six-month risk of second infection was three times higher among women with both, compared to neither, symptom (13.6 percent vs 38.3 percent). This suggests that it is the characteristics of the bacteria causing the initial infection and the extent of infection which predict recurrence.

Since I depended upon medical records for report of recurrence in a fairly mobile population, the risk of second UTI was probably underestimated. Moreover, sophisticated bacteriologic techniques were not used, making it impossible to determine how a particular bacteria and associated virulence factors might set the stage for recurrence. However, the results imply that future studies should examine bacterial and host characteristics as well as health behaviors to better understand the etiology of recurring UTI.

In conclusion, among this cohort of college women with initial UTI, demographic and behavioral factors did not distinguish between women who would and would not experience a second UTI during the six-month follow-up period. The strongest predictor of second infection was the presence of hematuria and urgency as symptoms of initial UTI. Future studies should use consider both behavioral and bacterial characteristics as potential determinants of recurring UTI.

ACKNOWLEDGMENTS

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Potential for Lyme Disease in Maine: Deer Survey of Distribution of *Ixodes dammini*, the Tick Vector

ROBERT P. SMITH, JR, MD, MPH, PETER W. RAND, MD, AND ELEANOR H. LACOMBE, BS, MT (ASCP)

Abstract: A survey of deer brought to tagging stations at 24 sites in Maine revealed the presence of the deer tick, *Ixodes dammini*, on 5.1 percent of deer. Ticks were found almost exclusively on deer from southwest coastal sites in the state. The potential for endemic Lyme disease in coastal Maine merits further study. (*Am J Public Health* 1990; 80:333–335.)

Introduction

The current geographic range of *Ixodes dammini*, the principal vector of Lyme disease in the eastern United States, is incompletely established. Continued spread of *I. dammini* in coastal New England accounts for new foci of Lyme disease in previously non-endemic areas.[1,2] Recent studies document the rapid rise of Lyme disease prevalence in some areas following colonization by the deer tick.[1] As public education remains the primary means of prevention of Lyme disease, knowledge of the regional distribution of the vector ticks enables targeted surveillance and focused public health interventions. Prior to 1985, there were no records of *Ixodes dammini* from Maine.* Although Maine has been considered a non-endemic area for Lyme disease, infected *I. dammini* were recently reported from one coastal Maine site[3] and *I. dammini* have been found on a deer shot in a northern inland site.[4] Three locally acquired seropositive cases of Lyme disease meeting Centers for Disease Control (CDC) criteria were reported to the Maine Bureau of Health between 1986 and 1988. We undertook a survey of ticks on deer brought to Maine tagging stations in 1988 to further describe the distribution of *I. dammini* at its northeastern limit in the United States.

Methods

Twenty-four deer tagging stations were selected for surveillance throughout Maine in consultation with state

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*Dearborn R: Personal communication. Department of Conservation, State of Maine.

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Inland Fisheries & Wildlife officials. Sites were chosen on the basis of high deer registrations during the previous hunting season, their proximity to representative coastal and inland habitat, and their accessibility to survey teams.

Volunteers including college biology students and their instructors, physicians, and public health professionals, and personnel from Inland Fisheries & Wildlife participated in two weekend surveys October 29 and November 5, 1988. Each team was supplied with a study protocol, an information sheet for hunters, data forms for each deer examined, 6 ml screw-top polyethylene specimen bottles with 70 percent alcohol, fine-point plastic tweezers, and rubber gloves. After obtaining permission from the hunter, each deer was examined for five minutes in a search for adult ticks that included the face, neck and ears of each deer. Ticks were removed with tweezers placed as close to the skin as possible. For each deer, the quantity and location of the ticks found, the times of the kill and of the examination, and the topography and vegetation of the kill site were recorded.

Type specimens of ticks were sent for species confirmation to Dr. J. E. Keirans, Department of Entomology, Smithsonian Institution, Washington, DC.

Results

Seventy volunteers visited 24 tagging stations at least one day of the hunting season. Two hundred and thirty-seven deer (141 male, 68 females, 28 sex unrecorded) taken in 66 towns, townships and plantations, including one island community, were examined. Twelve deer (5.1 percent) harbored *Ixodes dammini* ticks. A total of 33 ticks (all adults) were collected with a range of 1 to 11 ticks per infested deer. Ticks obtained from the hides of three freshly killed deer at a southwest coastal site are also included in the total. The average time of the examination of deer after the kill was 3.5 hrs (range, 0.5–46.5 hrs). Not included are time data from one island community where deer were examined one to three days after being tagged. Ten deer were found to harbor the "winter tick," *Dermacentor albipictus*. No other ectoparasites were obtained.

With the exception of one tick, all *I. dammini* were obtained from deer in the southwest coastal areas of Maine (see Figure 1). No ticks were found on deer from the one island studied. No association could be made between type of habitat at kill sites and deer infestation by *I. dammini* ticks.

Discussion

This study documents the presence of *Ixodes dammini* on deer from the southwest coast of Maine and a relative paucity of ticks from deer in inland areas. The habitat of southwest coastal Maine is similar to that in some other coastal New England areas with a high prevalence of Lyme disease, and it is characterized by deer herd densities as high as 20–50 per square mile. Although infection rates of ticks collected in this study were not determined, 12 percent of a small sample of adult and nymphal deer ticks from Acadia National Park harbored *Borellia burgdorferi*.[3] An infection rate of at least 10 percent has characterized other New England sites of *Ixodes dammini* colonization.[4,6] Although tick density on the deer examined was low relative to highly endemic sites,[7] the degree of risk for transmission of this zoonosis in southwest coastal Maine deserves further study. All three cases of endemic seropositive Lyme disease reported in Maine through 1988 were believed to have been acquired in this region. As there are no records of *I. dammini* or Lyme disease in Maine prior to 1985, we believe that *I. dammini* has appeared here recently.

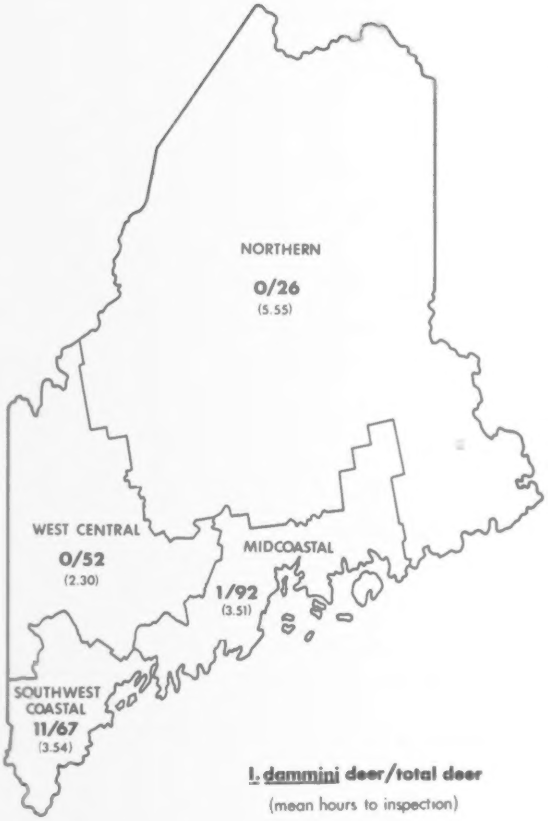

FIGURE 1—Number of Deer upon Which *I. dammini* Were Found versus Total Deer Examined in Four Geographical Areas in Maine

The sensitivity of deer surveys for the determination of presence of *Ixodes dammini* is unknown. The method of examination used, rigorousness of search, and training of personnel could alter its effectiveness. Reporting of other ticks on the deer (i.e., *D. albipictus*) is considered an indicator of the sensitivity of the search, but the expected prevalence of *D. albipictus* on northern New England deer is not known. A recent study of 370 deer in Vermont by trained wildlife personnel documented only two *Ixodes dammini* and no other ticks, although other ectoparasites were noted.** Comparative information of the effectiveness of deer surveys, dog seropositivity studies, and other methods of surveillance is needed. Our study is limited by marked differences in the number of deer available for examination at many tagging stations. The low number of deer examined in most inland sites and our inability to canvass all areas of the state systematically makes conclusions regarding the full range of *Ixodes dammini* in Maine premature. Nevertheless,

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**Mamolen M: Personal communication. Department of Health, State of Vermont.

the results of this deer survey coupled with the known ecology of southwest coastal Maine point to an emerging zoonosis in this area with potential public health importance.

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HRSA Health Care Program for the Homeless Reaches Its Legislative Goals

In a report to Congress on January 3, 1990, the Health Resources and Services Administration (HRSA) reported that its health care for the homeless program had been "successful in reaching its legislative goals." The report, which covered activities in calendar year 1988, was the first annual report on the program and noted the following:

- Within four months after signing of the authorizing legislation, the Steward D. McKinney Homeless Assistance Act, HRSA awarded $44.5 million in grants to 109 health care projects for the homeless in 43 states, the District of Columbia, and Puerto Rico.
- The 109 projects provided health care services to more than 231,000 homeless persons.
- Forty percent of the homeless persons served were children and their parents (living as a family unit) or homeless/runaway youths.
- Fifteen percent of those served were children ages 14 or younger.
- Thirteen percent of pregnant homeless females treated were teenagers.
- Sixty percent of persons served were minorities: 42 percent Black, 10 percent Hispanic, and 8 percent Native American, Asian, or other non-White.

The report concluded by saying, "The health care for the homeless program has demonstrated that the federal government, in concert with local communities, can bring health care services to homeless persons within a short period of time." For further information, contact Frank Sis, HRSA, US Public Health Service, DHHS, 5600 Fishers Lane, Room 14-43, Rockville, MD 20857. Tel: 301/443-3377.

* * * * * * *

Nursing Conference on Homelessness Scheduled May 3–4, 1990

A two-day forum for nurses and other health professionals involved in health care for the homeless projects will be held under sponsorship of a HRSA-supported project at State University of New York, Buffalo, May 3–4, 1990. Topics of the conference will include issues related to homeless children and mental health. For information about the conference, entitled "Homelessness: An Issue for the Nineties," contact: Continuing Nurse Education, School of Nursing, 14 Kimball Tower, SUNY at Buffalo, 3435 Main Street, Buffalo, NY 14214. Tel: 714/831-2391.

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Human Immunodeficiency Virus Seropositivity among Members of the Active Duty US Army 1985–89

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Abstract: Between October 1985 and June 1989, most active duty US Army soldiers were screened for human immunodeficiency virus (HIV) antibody. Of 648,032 screened soldiers in this analysis, 1,588 were HIV-antibody positive. In a multivariate analysis, correlates of positivity included: *age* [Adjusted Odds Ratios (ref <20 years) = 20–24 years, 3.7; 25–29, 9.3; 30–34, 15.7; ≥35, 15.9]; *being male*, [4.2]; *being Black or Hispanic (vs white)* [3.7 and 3.0, respectively]; *being single (vs married)* [3.8]; *assignment to an HIV endemic location* [1.7], and *having a medical occupation* [2.7, 2.7, and 2.6 for negligible, low, and high blood exposure professions, respectively]. Seropositivity rate ratios for medical vs non-medical personnel were 0.7 [95% CI = 0.4, 1.4] for females and 2.9 [95% CI = 2.5, 3.3] for males. For male medical personnel, being single (vs married) correlated strongly with antibody positivity [prevalence ratio = 3.4, 95% CI = 2.6, 4.6]. Excess HIV risk among medical personnel appeared largely attributable to factors other than occupational exposures. (*Am J Public Health* 1990; 80:405–410.)

Introduction

Effective programs to prevent the spread of human immunodeficiency virus (HIV) infection in the general population require the identification of target populations for intense health education and voluntary test-linked counseling.[1–4] Although much is known about the epidemiology of the acquired immunodeficiency syndrome (AIDS), the end stage of HIV infection, the value of this information for educational purposes is limited by the fact that AIDS cases generally reflect transmission risk factors that were operating years ago.[5,6] Prevention programs must reflect HIV infection-transmission patterns operating today and those likely to emerge in the near future. Prevalence and incidence surveys of generally healthy populations can help identify populations for targeted HIV prevention education.

Except for the national blood-donor screening programs, military HIV screening programs currently provide the only HIV seroprevalence data on large cross sections of the US population. Several papers have previously summarized the demographic risk factors for military recruit applicants, a group largely between the ages of 17 and 24.[7–9] In that population, the cumulative seroprevalence is 1.4/1,000 although over time the monthly rates have progressively dropped to below 1.0/1,000, a phenomenon that probably reflects self-deferral of potential applicants who either know from other sources or suspect that they are seropositive. Among recruit applicants, seropositivity is several-fold higher in Blacks, Hispanics, and males. Similar findings have been noted for members of the US Army Reserve Components.[10]

This paper, the largest detailed summary of military HIV demography published to date, expands on the recruit applicant data by focusing on active duty soldiers, a somewhat older group of persons who are more likely to be married and to have had substantial employment experience.

Methods

Population

On October 24, 1985 the Secretary of Defense directed that all 2.1 million active-duty US service members undergo mandatory HIV education and antibody screening.[11,12] This objective was addressed through a variety of programs that included mass screening of military units, screening prior to overseas deployment, blood donor testing, sexually transmitted disease (STD) clinic patient screening, case-contact tracing, and testing for clinical reasons. In 1988, the unit screening program was largely replaced by a birth month screening program designed to ensure that each soldier is tested at least once every 24 months.

HIV Testing

Sera collected as part of the mass screening program were tested with a licensed whole-virus lysate enzyme-linked immunosorbent assay (ELISA). This work was largely done at either a central civilian contract laboratory (Damon Laboratories, Irving, TX) for soldiers assigned to the continental US, Alaska, or Panama; at the 10th Medical Laboratory, Landstuhl, Germany for soldiers stationed in Europe; or at Tripler Army Medical Center for soldiers in Hawaii and the Far East. Positive samples were immediately retested in duplicate with the ELISA.[13] Repeatedly positive samples from all screening laboratories were tested for reactivity with HIV proteins by the Western (immune) blot method by the civilian contract laboratory.[14] From October 1985 until May 1987, specimens were classified as positive if antibodies to HIV-encoded protein GP41 and/or protein P24 plus P55 were detected.[15] In May 1987, the Association of State and Territorial Public Health Laboratory Directors criteria for a positive blot were adopted: at least two out of three of the major HIV protein bands P24, GP41, and GP120/160. Soldiers found positive on the initial Western blot had a second specimen drawn for verification purposes. A detailed Army-administered quality-assurance program characterized by monthly on-site laboratory inspections and specific performance criteria (including the successful analysis of both in-house controls and blind and unblinded panels) has been described elsewhere.[15]

Statistical Analysis

Because the Army population is undergoing constant turnover with only HIV-seronegative applicants allowed to enter and because during the period of data collection

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HIV-antibody positive soldiers have separated at a significantly faster rate than seronegative soldiers, this analysis, for reasons of clarity and reference, is confined to those soldiers on active duty when the program was just beginning, specifically on June 30, 1986. Laboratory results, social security number, first four letters of the last name, and test date were keyed onto computerized data files and merged with an abstracted copy of the June 30, 1986 active duty Army personnel file as supplied by the Defense Manpower Data Center, Monterey, CA. For soldiers with more than one test result recorded, only their most recent test was retained in the analysis data set. Antibody positivity was considered verified if an individual was reported by a separate clinical reporting system as HIV-infected (96.2 percent of those potentially considered positive were clinically verified as such) although in 22.9 percent of the cases clinical verification was based only on the documentation of an individual having had multiple positive HIV Western blots on separate specimens. Only soldiers verified to be HIV-infected prior to August 3, 1989 are considered HIV-infected for this analysis.

Seropositivity rates were calculated as both centrally screened and overall rates. The centrally screened rate is defined as the number of antibody positive soldiers who were reported as having been screened through the three centralized laboratories listed above divided by the total number of soldiers reported as screened by those laboratories.

Many soldiers found to be antibody positive were initially tested through testing activities conducted outside of the centralized testing laboratories (e.g., blood bank and some STD clinic screening). These individuals often were exempted from subsequent mandatory testing done by the centralized laboratories. The overall rate is based on a numerator that includes all antibody-positive active-duty soldiers who were in the Army on June 30, 1986 divided by the number centrally screened plus the number of positive soldiers not reported from the central laboratories. The centrally screened rate is a significant underestimate of the true rate because only 79 percent of the eligible antibody positive soldiers ever had a centralized screening test result reported. The overall rate is felt to be a modest overestimate because the denominator excludes a small though unquantifiable number of seronegative soldiers who either were tested but did not have a negative test reported or who were not centrally screened but were antibody negative in adjunct screening programs, such as blood bank and clinical screening. Some negative ELISA results from blood bank and clinical screening were not captured for epidemiologic purposes since not all of the adjunct screening was done by one of the three centralized HIV testing laboratories.

Because of centralized Western blot testing, positive blots from all testing programs were captured in an extremely complete fashion. For the most part, only overall rates and measures of effect based on overall rates will be described in the text since the overall rates are felt to best reflect the true rates. In the analyses of prevalence by occupation, broad occupational clusters were defined according to Army regulation, with minor modifications.[16] For the analysis of prevalence among medical personnel, the blood exposure groupings with minor modifications were those of Gardner, *et al.*[17]

To demonstrate the effect of a possible risk factor, seropositivity rate ratios with 95% confidence intervals were calculated.[18] In bar graphs, 95% CIs are based on the Poisson distribution. The main effects logistic regression analysis was performed with the Stepwise Logistic Regression (PLR) procedure of the BMDP Statistical Software Package (University of California at Los Angeles). All listed variables were forced into the model. Adjusted odds ratios were calculated by using logistic regression parameter estimates.[19] Confidence intervals for odds ratios and estimates of their statistical significance were derived from standard errors of the parameter estimates.[19]

Results

As of July 27, 1989, centrally screened results have been received for 647,700 soldiers out of the 763,675 individuals in the June 30, 1986 active-duty Army personnel file abstract. Demographically, this 84.8 percent sample (Table 1) constituted a representative sample of the whole Army. Three hundred and thirty-two soldiers who were verified as antibody positive but not reported as having been centrally screened were added to the analysis file. Of the 115,643 June 30, 1986 cohort soldiers without a documented HIV test result, 72 percent had separated from the active Army by December 30, 1988; 4 percent were accessions who had screened negative as part of the recruit-applicant screening instituted on October 1, 1985. The 28,863 remaining active-duty soldiers for whom valid test results were unavailable, were demographically similar to those tested based on race but women, married persons, officers and those at the extremes of the age spectrum appeared to be more likely not to have had a test reported.

Overall Seropositivity

The centrally screened seropositivity was 1.9/1,000 based on the 1,256 verified positive soldiers who had been centrally screened. The overall rate based on the 1,588 June 1986 cohort positives was 2.5/1,000. Of the 1,588 cohort soldiers who were verified as HIV antibody positive as of

TABLE 1—Characteristics of the Sample HIV Screened

Characteristic*	Army Population	Proportion of Army	Sample Population	Proportion of Sample
Age (years)		%		%
<20	78,543	10.3	66,430	10.3
20–24	282,792	37.0	234,073	36.1
25–29	169,579	22.2	147,229	22.7
30–34	108,248	14.2	96,530	14.9
≥35	123,665	16.2	102,846	15.9
Race/Ethnicity				
White	502,696	65.8	421,414	65.1
Black	205,192	26.9	177,991	27.5
Hispanic	27,047	3.5	23,635	3.7
Other	28,740	3.8	24,660	3.8
Sex				
Male	684,530	89.6	583,960	90.2
Female	79,093	10.4	63,691	9.8
Marital Status				
Single	313,655	41.1	260,183	40.2
Married	422,044	55.3	364,193	56.2
No Longer Married	27,001	3.5	22,593	3.5
Unknown	975	0.1	731	0.1
Rank				
Enlisted	652,490	85.4	551,974	85.2
Warrant	15,441	2.0	13,535	2.1
Commissioned	95,744	12.5	82,189	12.7
Length of Service (years)				
<1	115,513	15.1	98,209	15.2
1–3	267,969	35.1	219,130	33.8
4–6	110,967	14.5	96,922	15.0
7–9	83,915	11.0	74,232	11.5
≥10	185,311	24.3	159,161	24.6

*As recorded in June 30, 1986 military personnel file.

August 3, 1989 the reported sources of identification were: 809 (50.9 percent) through unit or birth month screening, 136 (8.6 percent) through clinical evaluations, 112 (7.1 percent) through pre-deployment screening, 101 (6.4 percent) through blood donor screening, 59 (3.7 percent) through STD clinic screening, 48 (3.0 percent) through self-referral, 17 (1.1 percent) through contact tracing, and 306 (19.3 percent) through other or unspecified sources.

Age

Antibody positivity varies significantly with age. For soldiers less than 20 years old at their most recent available test (a group heavily preselected to be negative by recruit-applicant testing), the overall rate was 0.5/1,000. Thereafter, there was a steady progression from an overall rate of 1.8/1,000 in those aged 20–24, to 3.1/1,000 in those 25–29 and to 3.4/1,000 in those 30–34. The rate among soldiers older than 34 drops to 2.2/1,000. A similar pattern was observed in the centrally screened rates.

Race or Ethnic Group

Soldiers who were Black or Hispanic had an overall seropositivity of 5.1/1,000 (900 cases) and 4.0/1,000 (95 cases), respectively. The HIV infection prevalence rate ratios compared to White soldiers were 4.0 [95% CI = 3.6, 4.4] for Blacks and 3.2 [95% CI = 2.5, 3.9] for Hispanics. Across age strata, Whites had a relatively flat prevalence trend; Blacks and Hispanics both showed a rise through the early thirties and then a fall (Figure 1). HIV infection would appear to be unevenly distributed among subsegments of the Hispanic population. The seropositivity rate for Hispanic soldiers with a home of record listed as Puerto Rico, New York, or New Jersey, a group assumed to be largely of Puerto Rican descent, was 5.6/1,000. In contrast, the Hispanic soldiers with a home of record in states along the Mexican border (California, Arizona, New Mexico, Texas) had a rate of 2.2/1,000 (PR = 2.5, 95% CI = 1.5, 4.3).

Sex

There were 1,516 verified positive cohort men, yielding an overall rate of 2.6/1,000. The corresponding female rate, based on 72 positive women, was 1.1/1,000. The crude male:female seropositivity rate ratio was 2.3 [95% CI = 1.8, 2.9]. Male:female overall rate ratios by race/ethnic group were 2.5 for Whites [95% CI = 1.5, 4.0], 3.0 for Blacks [95% CI = 2.3, 4.0], 2.1 for Hispanics [95% CI = 0.7, 6.7], and 3.1 for others (2 female cases). The overall rates in Black and Hispanic females were 1.9/1,000 (49 cases) and 1.9/1,000 (3 cases), respectively. The Black female rate exceeded the 1.3/1,000 overall rate in White males [RR = 1.4, 95% CI = 1.0, 1.9]. Age-specific male:female ratios did not show any definite trend across age strata (Table 2).

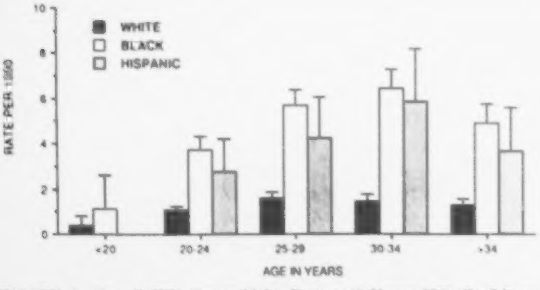

FIGURE 1—Overall HIV Seropositivity Rates and Upper 95% Confidence Interval Bounds for Active Duty US Army Personnel Stratified by Age and Race/Ethnic Group

Marital Status

Among the 1,588 cases, 639 were married, 878 were single, and 71 were no longer married, indicating that 44.7 percent of the cohort cases are or had been married based on the June 30, 1986 file. Compared to married soldiers, the HIV infection prevalence ratios for single soldiers and for those no longer married were 1.9 [95% CI = 1.7, 2.1] and 1. 8 [95% CI = 1.4, 2.3], respectively. Whereas overall rates were quite similar for married and single soldiers in the youngest age strata, single soldiers over 34 years of age had an overall prevalence rate that was 10.8 times that found in married soldiers in that stratum (Figure 2). Among White and Black soldiers, the rate ratios for single soldiers relative to married soldiers were 2.1 [95% CI = 1.8, 2.5] and 2.1 [95% CI = 1.8, 2.4], respectively. In contrast, the rate ratio for single Hispanic soldiers compared with their married counterparts was 1.1 [95% CI = 0.7, 1.7]. For males the overall single and married rates were 3.7/1,000 (844 cases) and 1.8/1,000 (607 cases), respectively. Among females the single (34 cases) and the married (32 cases) rates were both 1.1/1,000.

To explore the effects of age on estimates based on female marital status, female rates were directly age adjusted using the total male age distribution. The adjusted female rates were 1.2/1,000 for singles and 0.9/1,000 for marrieds.

Rank and Education

Based on 1,447 cases, the overall seropositivity rate for enlisted soldiers was 2.6/1,000. Warrant officers had an overall rate of 0.9/1,000 (12 cases) and commissioned officers an overall rate of 1.6/1,000 (129 cases). Within White, Black, and Hispanic groups, officer to enlisted prevalence ratios did not vary significantly (data available on request to author). Non-college graduates (less than 4 percent of soldiers in this cohort had not completed high school) had an overall rate of 2.6/1,000 (1,400 cases) in contrast to a rate of 1.9/1,000 (183 cases) for college graduates.

Occupational Category

HIV infection is distributed quite unevenly across job clusters (Table 3). Because health care workers had an elevated rate and because the potential for acquisition of infection in the course of performing job-related functions has been documented, several analyses were undertaken to better understand the source of risk for these individuals. Rate ratios comparing prevalence estimates for 42,697 health care providers [72.4% male] and 605,286 non-medical personnel [91.4% male] were 0.7 [95% CI = 0.4,1.4] for females and 2.9 [95% CI = 2.5, 3.3] for males. For male medical personnel being single [vs married] was a major correlate of

TABLE 2—US Army HIV Testing Overall Seropositivity Rate Estimates by Age and Sex

Age (years)	Male Rate* (No. Positive)	Female Rate* (No. Positive)	Male:Female Rate Ratio	95% CI
<20	0.4 (8)	0.9 (2)	0.5	0.1–2.2
20–24	1.9 (380)	0.7 (15)	2.8	1.7–4.7
25–29	3.4 (484)	1.2 (25)	2.8	1.9–4.1
30–34	3.5 (352)	2.0 (23)	1.8	1.2–2.7
≥35	2.3 (292)	0.9 (7)	2.6	1.2–5.5

*Rates are per 1000 persons

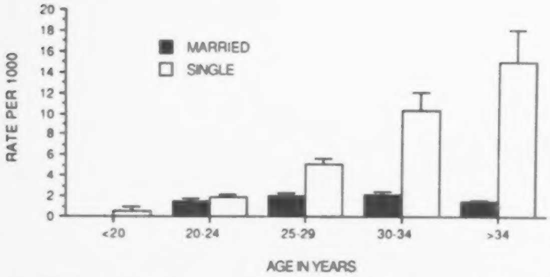

FIGURE 2—Overall HIV Seropositivity Rates and Upper 95% Confidence Interval Bounds for Active Duty US Army Personnel Stratified by Age and Marital Status

TABLE 3—US Army HIV Seropositivity Rates by Occupational Cluster

Job Cluster	Centrally Screened		Overall	
	Rate*	N Positives	Rate*	N Positives
Military Science	1.4	267	1.6	317
Electronics and Communications	1.9	140	2.6	189
Administration and Arts	3.8	212	4.9	270
Health Science	4.0	215	5.0	269
General Engineering and Science	1.7	36	2.1	44
Aviation	0.5	15	0.8	22
Intelligence/Military Police	1.0	42	1.3	58
Mechanical Maintenance	0.9	53	1.1	68
Missile Maintenance	1.3	11	1.4	12
Transportation, Supply, and Services	2.7	254	3.5	325
Other and Unknown	0.7	11	0.9	14

*Rate are per 1000 persons

positivity [prevalence rates = 3.4, 95% CI = 2.6, 4.6]. Among single medical males over the age of 34, the seropositivity ratio relative to their married counterparts was 15.9 [95% CI = 9.3, 27.3]. Rate ratios for male medical staff with high- and intermediate-level blood exposure relative to male care providers with minimal blood exposure were 0.9 [95% CI = 0.5, 1.5] and 1.1 [95% CI = 0.7, 1.6], respectively.

Multivariate Analysis (Adjusted Odds Ratios)

To control for the simultaneous effect of many mutually confounding factors, a main effects logistic regression model was constructed using the variables in Table 4. To reduce the preselection effect of recruit-applicant screening which started in October 1985, only soldiers with at least 10 months of service as of June 30, 1986 were included in the model. Noteworthy changes in odds ratios brought about by multivariate adjustment included the less obvious decline in risk for the oldest soldiers, the increased risk for men relative to women (adjusted OR = 4.2) and for singles relative to married soldiers (adjusted OR = 3.8). Adjustment significantly brought down the odds ratios associated with increasing length of service. The odds ratios by race/ethnic group, endemicity group, and medical blood exposure category were not significantly changed by the adjustment.

TABLE 4—Main Effects Multiple Logistic Regression Model for HIV Infection in the US Army

Variables	Crude Odds Ratio	Adjusted Odds Ratio	95% Confidence Interval
Age (vs <20)			
20–24	3.5	3.7	1.7–8.4
25–29	5.4	9.3	4.1–21.1
30–34	5.6	15.7	6.9–35.8
≥35	3.7	15.9	6.9–36.6
Sex (vs Female)	2.4	4.2	3.3–5.4
Race/Ethnic Group (vs White)			
Black	3.9	3.7	3.3–4.2
Hispanic	3.1	3.0	2.4–3.7
Others	1.8	1.6	1.2–2.2
Marital Status (vs Married)			
Single	2.3	3.8	3.4–4.4
Was Married	1.8	1.9	1.5–2.4
Endemicity*	1.7	1.7	1.4–1.9
Years of Service (vs <3 yrs.)			
4–6	1.1	0.8	0.7–1.0
7–9	1.4	0.8	0.7–1.0
≥10	0.9	0.6	0.5–0.7
Rank (vs Officer)	2.0	1.7	1.4–2.0
Blood Exposure Level (vs non-health care job)			
Negligible	2.2	2.7	1.8–4.0
Low	2.5	2.7	2.2–3.2
High	1.9	2.6	1.9–3.6

*Based on assignment in California, southern New York State, New Jersey, or the vicinity of the District of Columbia as of June 30, 1986.

Discussion

This analysis summarizes one of the largest population-based HIV antibody surveys undertaken to date. Although the observations are not directly generalizable to the 15 to 50 year old segment of the US general population due to various biases including the presumed underrepresentation of intravenous drug abusers and homosexuals in the military, the above results support the findings of other studies that HIV infection is an important issue for Black and Hispanic communities and for females.[7,20–23] The data appear to reasonably represent the people in the US Army on June 30, 1986. The entry and separation of over 100,000 people from the active component of the US Army each year plus the worldwide mobility of that population did not allow an HIV antibody screen to be completed on all members of the June 30, 1986 cohort.

The prevalence trend by age was similar to that noted for military recruit applicants and notable for a peaking in the late 20s and early 30s before tailing off in the mid-30s.[7] This trend in antibody positivity peaks several years before the peak described for AIDS cases. There are several possible explanations for the HIV antibody age pattern. First, those inclined to engage in risk behaviors may tend to moderate the intensity of their risk behaviors with age. Since the HIV virus has only been endemic in the US for an estimated 10–15 years, older soldiers may have passed through a period of more frequent risk taking before HIV was a significant presence. Second, older soldiers are more likely to be career soldiers than are those in the younger age strata. Soldiers who aspire to a military career and who have successfully adapted to the military environment (as suggested by many years of service) may be less inclined to practice some risk behaviors such as intravenous drug abuse and homsexuality because the practice of these behaviors is contrary to military policy. Since the early 1970s, random urine drug screens have been a regular occurrence in the Army, making even the covert use of illicit drugs a risk to a soldier's career. After adjusting for marital status, length of military service, and other variables in the multiple logistic regression mode), there is a steady increase in risk with age. Most importantly, however, these data suggest that significant risk of HIV infection begins in the late teens/early twenties and continues through the young adult period.

The observation that Black and Hispanic soldiers, even in the multivariate analysis, are at a three-to-four fold increased risk of being infected is consistent with civilian data.[7,20] The significant difference in prevalence trends by age for Blacks and Hispanics in contrast to Whites is again striking. A relatively flat prevalence trend for Whites and a dramatic rise and fall for Blacks and Hispanics may reflect a more rapid recognition of the HIV threat and its consequences on the part of Whites. The 2.5 fold increased risk of positivity for Hispanic soldiers with a home of record documented as Puerto Rico, New York, or New Jersey versus that for Hispanic soldiers with Mexican border state homes is reflective of Selik's finding that AIDS incidence appears to be higher in Hispanics of Puerto Rican descent than among other Hispanics.[24] Unfortunately, data on the risk behaviors directly responsible for the infections are not available to the authors. The observation that 50.4 percent of the 995 largely asymptomatic infected Black and Hispanic soldiers and 44.9 percent of the 536 largely asymptomatic White soldiers had at least seven years of service supports the impression that in the majority of cases the infection was acquired while the soldiers benefited from the economic, occupational, and social stability of a military environment. The fact that risk for both White and Black officers was similar to that of their enlisted counterparts further argues against a major role for current socioeconomic status in this military population.

As with all epidemics, the propagation of this epidemic depends on complex interactions among the agent, environment, and host. In the same way that high HIV prevalences (the agent-endemicity factor) among homosexual men tend to make sex acts (the behavioral-environment factor) more risky than similar acts among heterosexuals in a low HIV prevalence population, so too may risk activities in high prevalence minority populations lead to more infections than they might in lower prevalence communities. Although the data are inconclusive, race/ethnic differences in susceptibility (the host factor) to infection may be contributory factors also.[25-27]

In contrast to the 8:1 male:female AIDS case ratio in the US during the first half of 1989, the HIV infection prevalence ratio in these data is 2.4:1 (adjusted OR = 4.2:1 in multivariate analysis).[28] The prevalence ratio is similar to that reported for military recruit applicants.[7] In teenaged military recruit applicants (J. Brundage, unpublished data) and among people in the African HIV epidemic, the male-to-female ratio approximates 1:1.[21] Male:female prevalence ratios do not differ by race, but it is important to note that in the Army, the rate for Black and Hispanic women is 1.5 times higher than that found in White men, supporting the conclusion that women, particularly of racial/ethnic minorities, need special emphasis in prevention efforts. The marital status information further supports efforts to provide prevention education beyond single homosexual men. Although the marital status data do not allow any statement about the responsible risk activity, it shows that 45 percent of infected Army personnel are or were married. It would be reasonable to assume that most of these people participated in heterosexual sex at least some of the time and that many otherwise risk-avoiding heterosexuals were contacting the virus. One report of risk factors in a population of 57 US Army HIV-positive soldiers noted that 77 percent were heterosexually active, including 31 soldiers who said they were bisexual.[29] Centers for Disease Control AIDS data indicate that about 26 percent of reported AIDS cases are heterosexuals, although it is clear that most of these heterosexuals did not acquire their infections through sexual activity.[20]

Regarding risk by occupation, data presented here are consistent with other published estimates of the low risk for occupational acquisition of HIV infection in the medical setting.[30,31] The observations that excess risk for those in medical occupations appeared to be confined to males and that prevalence was not correlated with levels of occupational blood exposure are consistent with the conclusion that risk in Army health care personnel is largely due to factors other than occupational exposures. The logistic regression analysis also supports this and a comparable analysis of Army Reserve component data comes to a similar conclusion.[32]

The repeated testing of all active duty soldiers is allowing not only the ascertainment of antibody prevalence but also the current annual incidence of seroconversion. The first of these analyses for the period October 1985 to October 1987 identified 90 seroconvertors for a rate of 0.77 per 1,000 persons per year.[33] Current seroconvertors are carefully studied to ascertain potential shifts in transmission patterns since the design of prevention programs should reflect not only the transmission patterns in effect 10 or more years ago, but also recognize the role of emerging risk groups as the infection continues to spread, although probably more slowly, through initially unemphasized segments of the population.

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NCHS Survey: Characteristics of Persons Dying from Stroke

Cerebrovascular disease—or stroke—is the third leading cause of death in the United States, accounting for almost 150,000 deaths in 1986. The National Center for Health Statistics conducted a special survey based on a sample of such deaths among those 25 years of age and over which occurred in 1986. The survey analyzed the health and other care received in the last year of life and source of payment, health practices and behaviors, health status, and social and economic characteristics of those who died from stroke. These data will be used in disease prevention programs and to plan adequate health and support services for those with cerebrovascular disease. The major findings include:

- Nearly 70 percent of those who died from stroke were over the age of 75. Less than half of the deaths from all other causes occurred to those 75 years of age and over.
- More women than men died from stroke, 60 to 40 percent, respectively. Women were about three times more likely to be widowed at the time of death and only one-third as likely to be currently married.
- Most adults who died in 1986 had seen a physician five times or more in the year prior to death. Persons dying from stroke were more likely to have no physician visits or fewer than five visits; and to have resided in an institution during the year prior to death.
- Medicare was the primary source of payment for health care in the last year of life for most causes of death, including stroke.
- A larger proportion of those dying from stroke than from all other causes were reported to have hypertension.

For more information about this report, contact Sandra Smith, NCHS Public Affairs Officer, (301) 436-7135.

The Quality of Ambulatory Care in Medicare Health Maintenance Organizations

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Abstract: The quality of ambulatory care received by Medicare recipients who enrolled in health maintenance organizations (HMOs) was compared to the care received by fee-for-service (FFS) Medicare recipients, in a quasi-experimental, non-randomized design. Both samples were drawn from the four major geographic areas in the country, and included two types of HMO practices: staff/group models, and independent practice associations (IPAs). A panel of expert physicians developed criteria for evaluating ambulatory care, and medical record abstractions using these criteria were performed on 1,590 outpatient records: 777 FFS and 813 HMO (441 staff/group, 372 IPA). While individual items of medical histories and physical examinations were performed most often for staff/group HMO patients and least often in FFS patients, odds ratios (OR) for performance in staff/group HMO patients were particularly large for health maintenance items: tonometry (OR = 8.4), mammography (OR = 2.7), pelvic examination (OR = 5.3), rectal examination (OR = 2.9), fecal occult blood test (OR = 3.3). The results suggest that recommended elements of routine and preventive care are more likely to be performed for Medicare enrollees in staff/group HMOs than in FFS settings. (*Am J Public Health* 1990; 80:411–415.)

Introduction

Due to the rising costs of delivering health care to Medicare recipients, federal cost containment measures have been initiated. Among these initiatives, the use of prepaid care strategies has received increased attention. In 1982, the Health Care Financing Administration began a demonstration program of prepaid health care for Medicare recipients. The 27 participating health maintenance organizations (HMOs) provided comprehensive care, received prospective monthly payments for the enrollees, and assumed the financial risk of their complete health care. The purpose of the program was to show whether contracts with HMOs for Medicare recipients' health care could both reduce costs effectively and also maintain current standards of care. Since the demonstration program began, optional Medicare enrollment in HMOs has been extended, and there are now about one million Medicare enrollees in more than 100 separate HMOs.

Since HMOs reduce the cost of care mainly by replacing hospitalizations with less expensive ambulatory care, utilization rates of their enrollees have been observed to fall to as low as half the rates of Medicare recipients in traditional care arrangements.[1,2] Although the number of ambulatory visits in prepaid systems of care does rise to compensate for the reduction in hospitalizations, it is possible that such increases may strain the resources on which elderly HMO enrollees depend to manage their disabilities. Multiple chronic illnesses are the rule and many have significant functional limitations; outpatient management of elderly persons thus requires considerable time and effort. Hence, misgivings exist as to whether financial constraints, under prepaid systems like HMOs, may lead physicians to compromises that reduce the quality of outpatient care elderly persons need.[3,4]

To study how prepaid reimbursement affects the care Medicare patients receive in HMOs, the federal government funded the National Medicare Competition Evaluation. This comprehensive study examined not only the effectiveness of prepaid care in cutting Medicare costs, but also its effects on patient satisfaction with the care,[5] and whether HMO Medicare patient care matched the quality of that received by fee-for-service (FFS) Medicare patients. Several components of the evaluation are now either in press or in preparation. The evaluation of the process of routine and health maintenance care for ambulatory patients is presented here.

Methods

HMO Selection

There were several considerations in selecting the study's sample of HMOs. First, adequate representation of the four principal HMO models was a concern: staff HMOs, group HMOs, individual independent practice associations (IPAs), and network IPAs. Staff HMOs employ physicians and other providers directly; group HMOs contract one or more group practices. Individual IPAs contract with many individual physicians who care for HMO enrollees in their own offices; network IPAs contract not only with individuals but also with groups of physicians, with whom they usually share some of the risk for covered services. Because evidence suggests that IPAs function similarly to FFS practices, adequate representation of all four HMO models in the sample was important. Second, those plans with fewer than 2,000 Medicare enrollees were eliminated because they were unlikely to have enough cases for statistical power. Third, medical care often follows the conventional practices of the area, so adequate representation by geographic location was important. Finally, to secure the participation of HMOs and ensure confidence in the integrity of the study objectives, individual HMO plans were not identified in the analysis. From the pool of 27 HMOs in the demonstration, eight, distributed among the four major geographic areas, were selected for study. Four were staff or group model HMOs (staff/group) and four were IPA or network model HMOs (IPA).

Patient and Physician Selection

An important factor for selecting the sample of patients was that HMOs would have a high proportion of new evaluations. Routine care for new patients requires more

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detailed medical record-keeping than is required for care of established patients. Therefore, the samples of both HMO and FFS Medicare patients were limited to those who had new patient evaluations during the study time frame, the period from each plan's enrollment date (range: January 1983-May 1984) to March 31, 1986. From the HMOs selected for study, an appropriate sample of patients was identified by age, Medicare status, and the name of the primary care physician. Those enrolled in the HMO before becoming eligible for Medicare were excluded, as were rollovers, patients who had seen the same physician under FFS.

For the FFS sample, primary care physicians with eligible patients were identified and recruited using master files from the American Medical Association and the American Osteopathic Association. Within the geographic areas used in selecting HMOS, 590 primary care physicians were identified. They were recruited in random number order by telephone, and their primary provider status confirmed. Screening eliminated 342 physicians, of whom 42 percent were ineligible because they could not be reached by phone, and the others for a variety of reasons: currently practicing in a staff/group model HMO, not in an office-based practice, no new patients during the study time frame, not in primary care, dead or retired. Of the 248 eligible physicians, 135 refused to participate and 113 participated in the study (recruitment rate, 46 percent). About 58 percent of the participating primary care physicians in the FFS sample reported also seeing HMO patients not enrolled in a staff/group model HMO.

Each participating physician was promised confidentiality for the information collected in the abstracting process. Stipends were offered; they were accepted by 13 percent of participating physicians. Almost 60 percent of the participating physicians were board certified in family practice or internal medicine. About 50 percent had graduated from medical school by 1970. These data were not available for individual physicians in the HMO sample because agreement to participate in the study was reached with the HMOs. However, physician lists from HMO plans overall indicated professional profiles similar to those in the FFS sample. To avoid over- or under-representing a particular style of practice in the sample, records for about 10 patients were abstracted for each physician in the two groups.

Abstraction Instrument Design

An advisory panel of physicians was assembled to develop detailed criteria lists for evaluating ambulatory care. The physicians were chosen on the basis of experience in quality of care research, background with prepaid medical plans, and recommendations from colleagues. Among them were both general internists and clinicians with expertise in geriatrics.

The panel developed three categories of ambulatory patient evaluation: medical history-taking, physical examination, and preventive screening. The first two represent time-honored bedside techniques, and as such were employed to measure the quality of care in new patient evaluations. Elements of preventive care, such as immunizations and screening procedures, were emphasized because substantial evidence supports their efficacy in health maintenance for the elderly.[6] The abstraction instrument used is available by request.

Medical Record Abstraction

Data were collected by 10 registered nurse abstractors who had been specially trained. They first completed a set of exercises mailed to them along with the training manual, and then attended five days of closely supervised sessions. For periodic health examinations abstractors were explicitly instructed to follow current guidelines.[6] When a recommended examination was not performed, statements were accepted for deferral or noncompliance (e.g., history of a hysterectomy).

Interrater reliability was assessed by having all 10 abstractors and the principal investigator (SMR) review a medical record. Agreement among raters was assessed by a variation of the kappa statistic.[7] Kappa values between raters ranged from .45 to .85, indicating good to excellent agreement among nurse abstractors for most items. The lower interrater values occurred with medical-history items, and the higher values with physical examinations, screening tests, and procedures.

Analysis

While the study focused mainly on new patient evaluations, the follow-up time for the analysis of screening tests and procedures was extended from the initial patient evaluation to the time of the chart abstraction. Because there were differences between groups, the duration of time after the patient's initial evaluation was considered a potential confounder. Others were age and gender, since many screening items are both age- and gender-specific. To control for these confounders, a logistic regression model was used to test the relationship between the patient's membership in an IPA model HMO, a staff/group HMO or a FFS system of care, and the performance of individual health maintenance procedures. Performance of each health maintenance item was used as a single binary (0–1) dependent variable. Age was entered into the model as a continuous variable. Group status, gender, and duration of follow-up time were entered as categorical variables (duration of follow-up: 0 = <20 months, 1 = ≥20 months). Maximum-likelihood estimates were computed, and 95% confidence intervals were constructed around each odds ratio.[8,9] Since the purpose of the demonstration was to determine if the care received by HMO patients was at least equal to the care received by Medicare FFS patients, odds ratios were determined by comparing performance in each HMO group of patients to the performance in FFS patients.

Results

Records of 1,590 ambulatory patients were abstracted; of these, 777 were FFS and 813 were HMO patients. Of the 813 HMO enrollees, 441 (54 percent) were from staff/group HMOs and 372 (46 percent) were from IPAs. As shown in Table 1, the FFS and the two groups of HMO patients differed on several characteristics. Patients who had new patient evaluations from both types of HMOs had been in the practices longer than FFS patients. The staff/group HMO patients had more active health problems than either IPA or FFS patients; the FFS patients were most likely to have diastolic hypertension and obesity. All three groups had similar proportions of patients with diabetes mellitus; patients from IPA HMOs were somewhat more likely than the other two groups to have a history of a previous cardiovascular event. There were slight differences in race and living situations, with patients from both HMO groups more likely to be White and less likely to live alone; however, there were considerable missing data for these two variables.

When initial medical histories in the three groups were compared, as shown in Table 2, consistent differences were

TABLE 1—Patient Characteristics in the Three Plan Types

Characteristics	FFS n = 777 (%)	IPA n = 372 (%)	Staff/Group n = 441 (%)
Mean age at time of first visit to practice (SD)	73.6 (6.4)	73.0 (5.9)	72.5 (6.0)
Mean number of months between first and most recent visits to practice (SD)	17.5 (13.6)	22.2 (12.2)	22.6 (11.1)
Mean number of ongoing health problems (SD)	3.1 (2.0)	3.1 (2.0)	3.8 (2.8)
Number of female patients	496 (63.8)	195 (52.4)	243 (55.1)
Number of married patients	290 (51.2)	181 (57.3)	253 (61.7)
Number of patients with diastolic hypertension	434 (55.9)	208 (55.9)	300 (68.0)
Number of patients with diabetes mellitus	134 (17.3)	73 (19.6)	85 (19.3)
Number of patients with history of previous cardiovascular event	252 (32.4)	90 (24.2)	172 (39.0)

TABLE 2—Initial Patient Evaluation: Percentage of Records with Selected Medical History Elements, According to Plan Type

Historical Elements	FFS n = 777	IPA n = 372	Staff/Group n = 441
Record of past health problems or hospitalizations	84.0	87.4	89.1
Record of drug allergies	72.8	82.3	83.5
Listing of current medications	82.8	90.6	90.9
Explicit Review of systems:			
Cardiovascular	58.5	61.5	64.2
Gastrointestinal	52.1	56.9	59.5
Genitourinary	47.4	54.2	61.2
Mental status	38.5	40.2	51.7
Musculoskeletal	47.0	50.4	59.7
Neurologic	43.2	48.3	57.4
Respiratory	55.0	58.0	60.4
Record of history of tobacco	63.3	70.4	78.0
Record of history of alcohol	56.4	65.9	69.2
Record of social history	39.6	39.0	38.1
Record of functional status	55.3	49.0	68.3

TABLE 3—Initial Patient Evaluation: Percentage of Examination Elements and Laboratory Tests Recorded According to Plan Organization Type

Elements and Tests	FFS n = 777	IPA n = 372	Staff/Group n = 441
Examination of lungs	75.7	82.8	88.9
Examination of abdomen	74.1	84.7	88.4
Cardiac examination	73.6	80.1	90.9
Male genitourinary examination*	50.4	55.4	65.7
Peripheral vascular examination	50.2	49.5	59.4
Mental status examination	53.0	50.0	61.7
Neurologic examination	48.3	49.5	70.3
Urinalysis	66.4	70.2	88.9
Hemoglobin	79.4	79.8	93.9
Blood urea nitrogen or serum creatinine	71.3	75.0	93.2
Electrocardiogram	74.3	70.2	85.0

*Only males used in denominator

found. IPA and staff/group HMO patients were more likely to have their drug allergies and current medications listed. Staff/group HMO patients were more likely than either IPA or FFS patients to have a review of explicit system histories recorded, to have histories taken for tobacco and alcohol use, and to have a history of functional status recorded. Although IPA HMO patients had a greater likelihood of having an alcohol history taken than did FFS patients, they were less likely to have functional status information on their medical records.

Differences were also found among the groups with respect to the initial examinations, as shown in Table 3. As with medical histories, the consistent trend was for the records to show that examinations were more complete for patients in the two HMO models, especially the staff/group model.

In Table 4, performance of periodic health examinations is examined using logistic regression. The odds ratios with 95% CI comparing the two HMO groups to FFS patients, controlling for age, sex and duration of follow-up, are shown. Overall, HMO enrollees from both models were significantly more likely than FFS patients to have periodic health maintenance as part of new patient evaluations and follow-up; however, the enrollees in staff/group model HMOs were particularly likely to receive preventive care.

Discussion

The prospect of a continuing rise in the cost of health care for the elderly increases the likelihood of further initiatives to control it. Prepaid care has already been advocated as a promising means of reducing utilization. However, rationing Medicare recipients' health care by increasing the financial risk for the provider raises questions about the quality of care.[10,11] Our findings revealed that ambulatory HMO patients were likely to receive more of the routine evaluation and health maintenance procedures than were ambulatory FFS patients. Differences in performance of screening tests currently recommended for the elderly were particularly striking for enrollees in staff/group model HMOs.

The finding that patients in HMOs had higher performance rates of recommended health maintenance items than did FFS patients suggests that attention should be given to the effects of the structure of medical practice on quality of care. Several investigators have pointed out that physician

TABLE 4—Odds Ratios of Performance of Periodic Health Examinations Comparing Medicare Patients in HMOs and FFS with Logistic Regression Analysis*

Examination/Procedure	Staff HMO::FFS	IPA HMO::FFS
Visual Acuity	3.69 (2.83, 4.80)	1.47 (1.09, 1.97)
Tonometry	8.39 (5.41, 12.99)	1.92 (1.28, 2.87)
Breast Examination	2.00 (1.41, 2.82)	1.37 (.95, 1.99)
Mammography	2.67 (1.84, 3.88)	1.31 (.84, 2.04)
Pelvic Examination	5.26 (3.51, 7.89)	1.15 (.80, 1.66)
Rectal Examination	2.90 (1.89, 4.45)	1.17 (.77, 1.77)
Fecal Occult Blood Test	3.30 (2.54, 4.28)	.93 (.69, 1.24)
Influenza Immunization	1.03 (.76, 1.39)	.44 (.30, .66)

*95% confidence intervals in parentheses characteristics may play an important role, particularly in how they interact. Studies have suggested that medical practices incorporating significant peer interaction, such as HMOs, are more effective in reducing utilization than are those in which physicians practice independently, as in FFS practices.[12] It may also be the case that models with greater peer interaction more effectively support professional standards of care in the face of incentives to reduce utilization.[13,14] That feature of HMOs has the greatest potential to influence quality of care. Improved ambulatory care for Medicare recipients under prepaid systems may also be a result of HMOs' closer scrutiny of quality of care. Specific federal guidelines require quality assurance programs in HMOs.

While the results from this study may have important implications for health policy for the elderly, several limitations are worth noting. First, the study design did not allow for randomization of Medicare recipients into HMOs. Therefore, plan and physician selections may not have produced a representative sample, making application of the study results to the population of Medicare recipients in all HMO or FFS settings problematic; for example, it is possible that the FFS patients selected came from practices with unusually poor performance rates for standard elements of routine care. Nevertheless, among the elements of care received more often by the HMO enrollees, those constituting preventive care were most notable. Moreover, the performance rates for many of the preventive procedures received by patients in the study's FFS sample are comparable to rates reported in other FFS settings. For example, performance rates for FFS patients in nationwide surveys of clinical breast examination[15] and mammography[16] are remarkably similar to the performance rates among FFS patients in this study. Second, the study used a medical record review to determine performance rates for recommended items of routine care. It thus should be categorized as a study of the process of care and record-keeping. However, studies of the reliability and validity of medical records have established that technical aspects of medical care, such as preventive interventions, are likely to reflect actual performance.[17,18] Furthermore, for elderly patients the medical record itself is an especially important component of care, since their complex medical needs often call for multiple providers.

Although our study revealed that patients in HMOs receive medical care at least as competent as those in FFS practices receive, our analysis did not include an evaluation of individual HMOs. Our design plan was limited to representative models of group, staff, and two types of IPA HMOs, and only eight plans were studied. In that respect, the results from the evaluation of the demonstrations in this study should be interpreted with caution when establishing policy about all plans. Projects are currently under way to continue monitoring prepaid systems of care under Medicare.

Subject selection for this study was restricted to those who had been evaluated as new patients during the study time frame. However, the study did include data on patient follow-up visits for those periodic health examinations recommended after initial evaluation (e.g., annual mammography for women 50 to 59 years old). Therefore, the criteria for judging quality of care covered routine initial patient evaluations and follow-up periodic health examinations. The conclusions from the study are limited to those particular areas. Evaluations of the management of ambulatory patients with chronic illnesses, the follow-up of unexpectedly abnormal results, and the quality of care of inpatients are available in other reports.

The result that HMO enrollees received better preventive care may be due in part to the relative lack of reimbursement to FFS patients for the cost of screening tests. Others have noted that whether a patient undertakes recommended mammography, for example, is directly related to whether or not she can afford it.[19] Previous research has shown that preventive health care utilization is higher among HMO enrollees than among non-enrollees.[20] However, aside from procedures affected by reimbursement, HMO patients in this study were also more likely than FFS patients to have other examinations specifically recommended for the routine care of the elderly. For example, HMO enrollees were more likely to have current medications listed, an item especially important for elderly patients,[21] and enrollees in staff/group models were more likely to have functional status assessments. Furthermore, intraocular pressure and visual acuity were screened much more often in HMO enrollees than in FFS patients.

Efforts to regulate Medicare health costs must undergo careful scrutiny mindful of elderly patients' extensive medical needs. In this study, the scope of the conclusions is limited in that the quasi-experimental design restricted the findings to new patient evaluations in a demonstration program. The data showed that the elderly enrolled in HMOs received ambulatory care at least equal to what patients in FFS settings received; however, further evaluations are needed to verify that conclusion.

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$5 Million Available for Pediatric AIDS Grants

Approximately $5 million is available for competitive grants for three types of AIDS (acquired immunodeficiency syndrome) Projects, according to an announcement from the federal Health Resources and Services Administration (HRSA). Six to eight awards will be made for **Pediatric AIDS Health Care Demonstration Projects** intended to demonstrate or develop:

- effective ways to prevent infection; particularly through reduction of perinatal transmission;
- community-based family-centered coordinated services for infected infants, children, and youth;
- programs to reduce the spread of HIV (human immunodeficiency virus) infection to vulnerable populations of young people.

Up to three grants will be made for **Pediatric AIDS Comprehensive Centers** intended to:

- establish public-private consortia among institutions and service providers; and
- develop a service, training, and research capability to support the community service base.

Four to six grants will be awarded for **National Issues of High Priority in Pediatric AIDS**. These grants will address critical issues for people with AIDS and their families, especially in prevention, early detection, and service provision. One grant in this category will establish a National Resource Center in Pediatric AIDS to identify and deal with issues that prevent patients and families from receiving services.

Public and private entities, nonprofit and for-profit, are eligible to apply for the grants. The *application deadline is April 9*. These AIDS project grants were announced in the February 13 issue of *Federal Register*.

Application materials are available from: Grants Management Officer, Bureau of Maternal and Child Health and Resources Development, Suite 100A, 12300 Twinbrook Parkway, Rockville, MD 20852; (301) 443-1440.

AIDS Risk Behavior Patterns among Gay Men in Small Southern Cities

JEFFREY A. KELLY, PHD, JANET S. ST. LAWRENCE, PHD, TED L. BRASFIELD, RTT, L. YVONNE STEVENSON, MS, YOLANDA E. DIAZ, MSW, AND ALLAN C. HAUTH, MA

Abstract: Men entering gay bars in three small southern cities were administered questionnaires to assess AIDS (acquired immunodeficiency syndrome) risk knowledge, perceived norms concerning the social acceptability of AIDS risk precautions, and personal sexual behavior. Seventy-seven percent (N = 355) of male patrons completed the measures. While risk knowledge levels were high, respondents did not perceive strong norms favoring risk reduction precautions. Rates of risk behavior were considerably higher than those reported for gay men in large urban epicenters. (*Am J Public Health* 1990; 80:416–418.)

Introduction

A number of cohort studies provide evidence that gay men in large-city AIDS (acquired immunodeficiency syndrome) epicenters have substantially modified behavior patterns to reduce risk for HIV (human immunodeficiency virus) infection.[1–5] Recent studies suggest that fewer than 20 percent of gay men in New York City and under 10 percent of gay men in San Francisco have engaged in unprotected anal intercourse during the past 12 months.[6,7] As a result, the prevalence of HIV infection among these populations has essentially reached a plateau, although at very high saturation levels.[8,9]

Less is known about the risk behavior patterns of homosexual men in smaller cities where HIV infection prevalence is still relatively low. For AIDS primary prevention efforts to be successful in these areas, there must be a rapid shift in sexual behavior norms and practices comparable to those already achieved by most gay men in AIDS epicenters. Limited research to date suggests that has not occurred.[10–12] The current study examined AIDS risk knowledge, perceived sexual behavior norms, and risk behavior patterns in a sample of homosexual men in small cities with currently low prevalence of AIDS.

Methods

Over consecutive three-night periods in the spring of 1989, all males entering gay clubs in three southern cities (Monroe, LA; Hattiesburg, MS; and Biloxi, MS) completed written questionnaires. A mean of 77 percent (range 56 percent to 85 percent) of all men entering each bar completed the study measures individually and without personal identifiers (N = 355). The three cities have metropolitan populations of 50,000 to 56,600 persons and are located at least 60 miles from any larger population center. While data on the number of AIDS cases in these cities are not available, the annual AIDS incidence in the study states (8.8/100,000 population in Louisiana and 5.8/100,000 in Mississippi) is much lower than that of states which have original AIDS epicenters (36.9/100,000 in New York and 20.9/100,000 in California).[13] Two of the cities have one gay bar while the other has two clubs; all were surveyed.

Each questionnaire requested demographic information and included probe measures of AIDS risk behavior knowledge, perception of norms concerning the acceptability of taking risk reduction precautions, and self-reported sexual activities over the preceding two months. The AIDS risk behavior knowledge measure consisted of 23 true/false items which tap practical understanding of high-risk sexual practices and risk reduction steps.[10] Five Likert-scale items (from 1 = strongly agree to 5 = strongly disagree) assessed the respondent's perception of peer norms and social consequences of taking risk reduction steps (Table 1). Each respondent also reported the number of times and the number of different sexual partners over the past two months for the sexual activities listed in Table 2.

Results

Preliminary analyses revealed no significant differences in risk behavior rates, knowledge test, and social norm scores across the three cities so data from all cities were pooled. The mean age of respondents was 28.3 years (SD = 7.9), mean educational level was 14.6 years (SD = 2.7), and 91 percent of the sample was White.

On the knowledge test, subjects correctly answered an average of 20.8 (SD = 2.3) of 23 items indicating accurate understanding of risk behavior. However, the social norms measure reflected only mixed perception that friends behaved safely or that insistence on condom use and safer sex practices would be well-accepted by others. Table 1 presents the distribution of responses for each item. While 81 percent of respondents were neutral or disagreed with the statement that condom use implies a lack of trust in a sexual partner, only 31 percent agreed or strongly agreed that their friends always use condoms during intercourse, only 43 percent agreed that peers used condoms or engaged in non-insertive practices, and 54 percent believed friends talk about safer sex more than they practice it. Respondent's mean score on these questions was 16.0 (SD = 4.1) of a maximum possible score of 25.

Eighty-three percent of respondents reported sexual activity over the preceding two months; 52 percent reported one partner and 48 percent reported multiple sexual partners. As Table 2 shows, 25 percent of all respondents reported engaging in unprotected insertive, and 23 percent unprotected receptive, anal intercourse during the past two months. For the total sample, the mean frequencies of these behaviors were 1.41 and 1.04 occurrences respectively. However, for subjects who engaged in any unprotected anal intercourse, the mean frequencies were 5.55 and 4.59 respec-

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TABLE 1—Responses to Items Assessing Perception of Social Norms Concerning AIDS Risk Precautions

	Respondents Who:				
Statements	Strongly Disagree N (%)	Disagree N (%)	Neutral N (%)	Agree N (%)	Strongly Agree N (%)
My friends always use condoms during intercourse	62 (18)	52 (15)	129 (36)	61 (17)	50 (14)
My friends talk about safer sex more than they actually practice it	35 (10)	36 (10)	92 (26)	85 (24)	107 (30)
My friends believe that insisting on safer sex implies that you don't trust your partner	146 (41)	40 (11)	102 (29)	36 (10)	31 (9)
Safer sex is completely accepted by my friends	27 (8)	37 (10)	85 (24)	74 (21)	132 (37)
My friends either use condoms or have sex that doesn't involve fluid exchange	47 (13)	44 (12)	114 (32)	63 (18)	87 (25)

N = 355

TABLE 2—Risk Behavior Patterns Reported by Study Respondents

Behavior	Number and Percentage of Total Sample Reporting Behavior in Past Two Months	Mean Frequency for Entire Sample in Past Two Months	Mean Frequency for Respondents Who Report Any Occurrence of the Behavior in Past Two Months	Mean Number of Partners for Behavior in Past Two Months for Subjects Who Report Any Occurrence of the Behavior
Unprotected anal intercourse as insertive partner	88 (25)	1.35 (4.85)	5.55 (9.45)	1.11 (.56)
Unprotected anal intercourse as receptive partner	81 (23)	1.11 (4.06)	4.59 (7.08)	1.03 (.62)
Unprotected oral intercourse as insertive partner, fluid exchanged	123 (35)	1.59 (3.76)	5.46 (11.12)	1.58 (2.91)
Unprotected oral intercourse as receptive partner, fluid exchanged	102 (29)	1.85 (6.11)	6.67 (11.08)	2.19 (5.75)

N = 352

tively. A proportionate calculation revealed that condoms were used by the insertive partner in only 51 percent of all anal intercourse occasions. In addition, 35 percent of the sample reported unprotected insertive, and 29 percent in receptive, oral intercourse with fluid exchange over the same two-month period. Finally, a median split was performed to subdivide the sample into those with high and those with low total scores on the social norm measure. Respondents with high social norm scores engaged less often in insertive and receptive unprotected anal intercourse and in receptive oral sex (Table 3).

TABLE 3—Means and 95% Confidence Intervals Comparing Sexual Behavior Practices of Respondents Scoring Above and Below Median in Social Norm Total Score

	Perceived Social Norm Score Median Split	
	Respondents in Upper Median	Respondents in Lower Median
Sexual Practice Variable	M (95% CI)	M (95% CI)
Unprotected Insertive Anal Intercourse	0.6 (0.2,0.9)	2.3 (1.2,3.4)
Unprotected Receptive Anal Intercourse	0.4 (0.1,0.7)	1.7 (0.9,2.5)
Unprotected Insertive Oral Intercourse, Fluid Exchanged	1.3 (0.6,2.0)	2.6 (1.2,4.0)
Unprotected Receptive Oral Intercourse, Fluid Exchanged	0.8 (0.3,1.4)	3.0 (1.7,2.7)

N = 352

Discussion

Rates of high-risk sexual behavior in these cities are much greater than those recently reported for gay men in most large AIDS epicenters. The percentage of respondents in this study reporting unprotected anal intercourse during just the past two months is two to three times higher than the percentage of San Francisco gay men who report engaging in this practice during the last 12 months.[7] Conclusions that homosexual men have already effectively made changes needed to reduce risk for new HIV infection based upon large-city cohorts may not generalize to smaller cities.

Factual knowledge about AIDS was high in this sample. However, we suspect that gay men in small cities encounter fewer social supports for making AIDS precautionary behavior changes, do not benefit from well-organized community social structures which permit efficient behavioral norm redefinition, and still perceive AIDS as a distant threat. These social norms have considerable impact on gay mens' sexual behavior.[14]

The cross-sectional method used in this study does not permit conclusions to be drawn concerning behavior changes which may have occurred from earlier time periods, the sample is limited to patrons of gay bars in several southern cities, and there is a wide range in behavior reported by these respondents. Nonetheless, intensified primary prevention interventions are necessary for the many gay men in communities outside the urban areas who remain at great risk for HIV infection.

ACKNOWLEDGMENTS

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NIH Consensus Development Conference on Colon/Rectum Cancer, April 16–18

Adjuvant Therapy for Patients with Colon and Rectum Cancer is the subject of an upcoming consensus development conference sponsored by the National Institutes of Health (NIH). The conference, scheduled for April 16–18, 1990, in Masur Auditorium of the NIH Clinical Center, is open to the public.

The purpose of the conference is to reach agreement on adjuvant therapy for patients with colon and rectum cancer, as well as directions for future research.

Large bowel adenocarcinoma is a major public health problem in the United States. More than 150,000 new cases of colon and rectum cancer will be diagnosed in 1990 and approximately 75 percent will have a primary surgical resection with the hope of complete tumor eradication. Despite this high resectability rate, nearly half of all colorectal cancer patients die of metastatic tumor.

The conference will bring together specialists in surgery, oncology, gastroenterology, pathology, statistics, and other relevant fields. On the first two days, experts will present current scientific thinking about the diagnosis, management, and prevention of adjuvant therapy for patients with colon and rectum cancer, and concerned voluntary organizations will be invited to make statements. On the third day, after considering the scientific evidence, the consensus panel will present its draft report and invite comments from the audience. Glenn D. Steele, Jr., MD, PhD, New England Deaconess Hospital, will chair the panel.

To register for the Consensus Development Conference on Adjuvant Therapy for Patients with Colon and Rectum Cancer, or to obtain further information, contact: Conference Registrar, Prospect Associates, Suite 500, 1801 Rockville Pike, Rockville, MD 20852; (301) 468-MEET. FAX: (301) 770-5164.

HIV-1 Counseling and Testing Sites, Minnesota: Analysis of Trends in Client Charactertistics

Richard N. Danila, PhD, MPH, James M. Shultz, MS, PhD, Michael T. Osterholm, PhD, MPH,
Keith Henry, MD, Margaret L. Simpson, MD, and Kristine L. MacDonald, MD

Abstract: We report here a summary of the data obtained from two HIV-1 antibody counseling and testing sites in Minneapolis-St. Paul for the first 48 months of operation (24,911 persons tested). The HIV-1 antibody seroprevalence rate for all persons tested was 5 percent. The highest seroprevalence rates were in male homosexual/bisexual intravenous drug users (23 percent) and homosexual/bisexual men (13 percent). There was a significant decrease in the HIV-1 antibody seroprevalence rate among clients during the 48-month period from 14 percent in the first six months to 3 percent in the last six months. This decrease coincided with an increase in the number of low-risk female clients and low-risk heterosexual male clients, and a decrease in the number of homosexual/bisexual males participating in the programs. These findings suggest the need for development and implementation of other strategies to identify and reach persons at highest risk for HIV-1 infection. (*Am J Public Health* 1990; 80:419–422.)

Introduction

In July 1985, the Minnesota Department of Health (MDH) established human immunodeficiency virus (HIV-1) counseling and testing sites at the two major public health sexually transmitted disease clinics in the Twin Cities (Minneapolis, St. Paul) to allow persons at risk for HIV-1 infection to ascertain their HIV-1 antibody status.

To examine the potential role of these Minnesota programs in HIV-1 infection prevention efforts, data on client demographic and risk factor characteristics and HIV-1 antibody seroprevalence were analyzed for the first 48 months of operation (July 1985 to June 1989).

Methods

The two programs provide free services to all persons requesting them. Prospective program clients are scheduled for appointments by telephone. All clients are interviewed by trained nurses or clinicians according to protocols established by the US Public Health Service Centers for Disease Control[1] and the MDH (available upon request to authors). During pre-test and post-test sessions, staff assess client risks for HIV-1 infection, interpret HIV-1 antibody test results, and provide individualized risk reduction and disease prevention counseling. Self-reported demographic and risk factor data are recorded on standardized client information sheets.

Serum specimens are initially screened for HIV-1 antibody using enzyme immunoassay (EIA) test kits from Abbott Laboratories, Electronucleonics, or Genetic Systems. Specimens are considered reactive if the absorbance values are above the cutoff levels specified by the manufacturers' instructions. Initially reactive specimens are tested a second time; specimens that are repeatedly reactive are then tested using the Western blot technique.[2] Western blot testing utilizes HIV-1 antigen from Bio-tech Dupont, Rockville, Maryland, or Litton Bionetics, Charleston, South Carolina. Specimens are considered positive by Western blot assay if the following bands are detectable: p24 or p31 and either gp41 or gp120/gp160, with or without additional HIV-1 characteristic bands.[3] Specimens with other HIV-1 band patterns not meeting the criteria for a positive assay are considered to be indeterminate.

Data were analyzed for 24,911 clients who had HIV-1 antibody serology completed from July 1985 through June 1989. Male clients were categorized into one of five HIV-1 transmission categories: homosexual/bisexual, homosexual/bisexual intravenous drug user, heterosexual intravenous drug user, heterosexual sex partner of a person at high risk for HIV-1 infection, or other (immigrant from a high-risk country or no identified risk). Female clients were categorized into one of three HIV-1 transmission categories: intravenous drug user, sex partner of a person at high risk for HIV-1 infection, or other (immigrant from a high-risk country, female prostitute, or no identified risk). Data were analyzed using standard univariate methods and chi-square test for trend measures.[4]

Results

Between July 1, 1985 and June 30, 1989, 25,306 client visits were recorded; 24,911 clients (98 percent) were tested for HIV-1 antibody and 395 clients (2 percent) declined testing after the pre-test visit. An increase occurred in the number of clients tested per month over the first 36 months with a leveling over the last 12 months (Figure 1). The HIV-1 antibody seroprevalence rate for all clients tested was five percent (1,299 of 24,911).

First-time and Repeat-Test Clients

For tested clients, 18,288 (73 percent) were persons being tested for HIV-1 antibody for the first time (first-time clients) and 6,623 (27 percent) were clients who had been tested at least once before (repeat-test clients). The HIV-1 antibody seroprevalence rate for first-time clients (898 of 18,288; 5 percent) was lower than the rate for repeat-test clients (401 of 6,623; 6 percent). For the repeat-test clients, 2,672 (40 percent) had been tested previously at a health department site, 353 (5 percent) at a family planning clinic, 306 (5 percent) at a physician's office, 81 (1 percent) at a plasma center, 58 (1 percent) at a blood collection facility, and 3,153 (48 percent) at other unspecified sites. Of the 401 clients who were HIV-1 antibody positive on the repeat test: 175 (44 percent) had tested positive for HIV-1 antibody previously; nine (2 percent) had prior tests with indetermi-

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From the Acute Disease Epidemiology Section, Minnesota Department of Health, Minneapolis (Danila, Shultz, Osterholm, MacDonald); St. Paul-Ramsey Medical Center, and St. Paul Division of Public Health (Henry); and Hennepin County Medical Center, Minneapolis (Simpson). Address reprint requests to Richard N. Danila, PhD, MPH, Acute Disease Epidemiology Section, Minnesota Department of Health, 717 Delaware Street, SE, Minneapolis, MN 55440. This paper, submitted to the Journal July 3, 1989, was revised and accepted for publication October 31, 1989.

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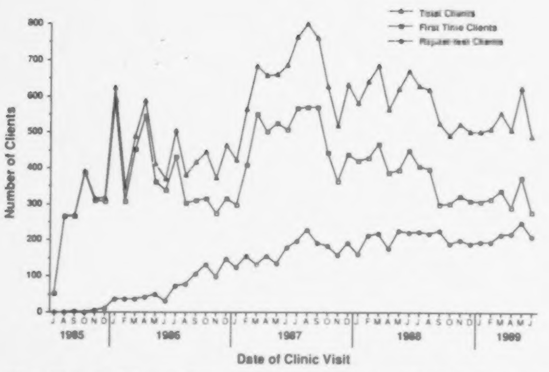

FIGURE 1—Number of Clients Attending Counseling and Test Site Programs Tested for HIV-1 Antibody by Visit. Status by Month, Minnesota, 1985-89

nate Western blot test results; 107 (27 percent) had previously tested either EIA negative or EIA reactive and Western blot negative; 33 (8 percent) had tested EIA reactive previously but had not had confirmatory Western blot testing performed; and information on previous testing was unavailable for 77 (19 percent).

Most clients (91 percent for first-time clients) were between 20 and 49 years of age. The HIV-1 antibody seroprevalence rate was higher in males than females (Table 1). The HIV-1 seroprevalence rate for Blacks was 1.6 times higher than the rate for Whites and the rate for Hispanic clients was 1.8 times as high as the rate for Whites.

The highest HIV-1 seroprevalence rate was found in male homosexual/bisexual intravenous drug users (23 percent; Table 2). Rates were also analyzed by transmission category and race for first-time clients; higher rates were consistently found in Black and Hispanic males (data available upon request to authors). HIV-1 seroprevalence rates for first-time clients were 6 percent and 4 percent at each of the two clinics, respectively; the higher seroprevalence rate was due to a larger proportion of homosexual/bisexual male clients seen at that site.

TABLE 1—HIV-1 Antibody Seroprevalence Rates by Gender, Age, and Race for First-Time Clients; Minnesota Counseling and Testing Sites; July 1985 through June 1989

Client Characteristics	No. Tested (%)	No. HIV-1 Anitbody Postive	Seroprevalence Rate
Total	18,288 (100.0)	898	4.9
Gender			
Male	12,085 (66.1)	876	7.2
Female	6,128 (33.5)	22	0.4
Unrecorded	75 (0.4)	0	0.0
Age Group (years)			
0–12	22 (0.1)	2	9.1
13–19	612 (3.3)	9	1.5
20–29	7,495 (41.0)	384	5.1
30–39	6,830 (37.3)	377	5.5
40–49	2,307 (12.6)	89	3.9
50–59	598 (3.3)	22	3.7
60 and over	229 (1.3)	3	1.3
Unrecorded	195 (1.1)	12	6.2
Race			
White	16,609 (90.8)	788	4.7
Black	1,020 (5.6)	79	7.7
Hispanic	188 (1.0)	16	8.5
Other*	303 (1.7)	11	3.6
Unrecorded	168 (0.9)	4	2.4

*Includes Native American and Asian/Pacific Islander.

TABLE 2—HIV-1 Antibody Seroprevalence Rates by Client Transmission Categories for First-Time Clients, Minnesota Counseling and Testing Sites, July 1985 through June 1989

HIV-1 Transmission Categories	No. Tested (%)	No. HIV-1 Antibody Positive	Seroprevalance Rate (%)
Males			
Homosexual/Bisexual	5,866 (48.5)	761	13.0
Homosexual/Bisexual Intravenous Drug User	172 (1.4)	40	23.3
Heterosexual/Intravenous Drug User	803 (6.7)	50	6.2
Heterosexual Sex Partner at High-Risk	1,403 (11.6)	5	0.4
Other Heterosexual/ No Identified Risk	3,841 (31.8)	20	0.5
Total	12,085 (100.0)	876	8.1
Females			
Intravenous Drug User	405 (6.6)	6	1.5
Sex Partner at High-Risk	1,881 (30.7)	7	0.4
Other/No Identified Risk	3,842 (62.7)	9	0.2
Total	6,128 (100.0)	22	0.4

Analyses of Trends Over Time

A significant downward trend in the monthly HIV-1 antibody seroprevalence rate was noted over time for all clients tested (both first-time and repeat-test clients) ($p < 0.001$) (Figure 2). The seroprevalence rate for all clients tested during the six months of testing in 1985 was 14.3 percent. The seroprevalence rate dropped to 8.3 percent during 1986 and declined further to 3.9 percent during 1987 and to 3.5 percent during 1988. During the first six months of 1989, the overall seroprevalence rate was only 2.5 percent.

The number and proportion of repeat-test clients increased significantly over time ($p < 0.001$) (Figure 1). Repeat-test clients accounted for 1.2 percent of all clients during the first six months of operation in 1985, but comprised 15.9 percent of all clients in 1986, 26.1 percent of all clients in 1987, 35.0 percent of all clients in 1988, and 40.2 percent of all clients during the first six months of 1989. This increase in the proportion of repeat-test clients did not account for the observed decline in the overall HIV-1 antibody seroprevalence, since the seroprevalence rate for repeat-test clients was actually higher than for first-time clients, as noted previously.

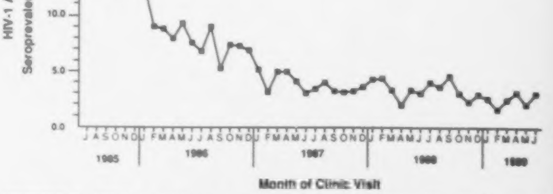

FIGURE 2—HIV-1 Antibody Seroprevalence Rates among Clients Attending Counseling and Test Site Programs by Month, Minnesota, 1985–89

Analysis of data from first-time clients only demonstrated that the HIV-1 seroprevalence declined from 14.2 percent for the first six months of operation in 1985 to 1.8 percent for the last six months of operation in 1989. The HIV-1 seroprevalence rate for repeat-test clients similarly declined from 26.3 percent in the first six months in 1985 to 3.5 percent in the last six months of operation in 1989 (data available on request to authors).

Major changes were noted in client characteristics over the 48-month period. Initially, most clients were at increased risk for HIV-1 infection; however, the number and proportion of clients at low risk for HIV-1 infection increased over time. During the first six months of the program, females comprised 8 percent of all first-time clients, whereas they made up 43 percent of all first-time clients during its last six months. During the first six months, 54 percent (69 of 128) of the females were at increased risk (intravenous drug user or sex partner of a person at high risk) and 46 percent (59 of 128) were at low risk; whereas, during the last six months, 33 percent (261 of 795) were at increased risk and 67 percent (534 of 795) were at low risk. There were no seropositive females in the first six months and only one for the last six months. During the first six months, 81 percent of all first-time clients were homosexual/bisexual males and 10 percent were heterosexual males; whereas, only 14 percent were homosexual/bisexual males and 43 percent were heterosexual males during the last six months of observation (Figure 3). For homosexual/bisexual male clients, a significant decline in HIV-1 antibody prevalence over time was noted ($p < 0.05$) (Figure 4).

Discussion

Counseling and testing site programs were originally conceived, funded, and opened in 1985 to assure safety of the blood supply by providing alternate sites for potential volunteer blood donors to ascertain their HIV-1 antibody status.[5] They have become a major cornerstone for AIDS (acquired immunodeficiency syndrome) prevention programs throughout the country and an integral part of the Minnesota statewide plan for AIDS prevention.[6] In evaluating the contribution of these programs to primary prevention of HIV-1 transmission, three issues need to be considered.

- First, are they currently serving clients most at risk for HIV-1 infection?
- Second, are they successful in HIV-1 infection prevention efforts through effecting and maintaining behavior change in clients?

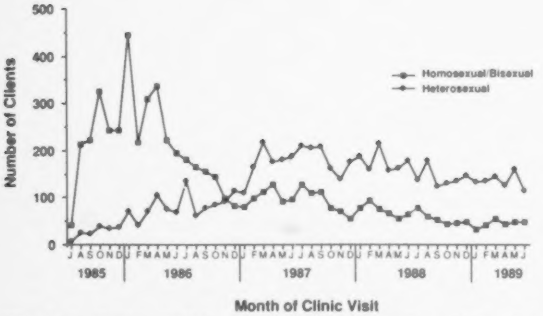

FIGURE 3—Number of Male Clients Attending Counseling and Test Site Programs by Month, by Sexual Orientation, Minnesota, 1985–89

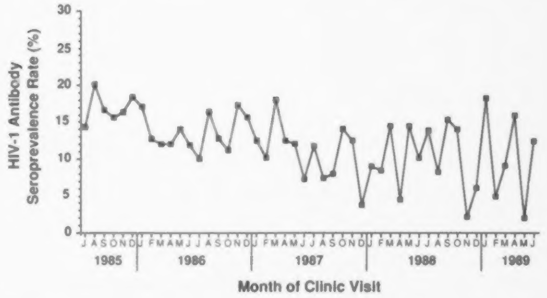

FIGURE 4—HIV-1 Antibody Seroprevalence Rate for Homosexual/Bisexual Males Attending Counseling and Test Site Programs by Month, Minnesota, 1985–89

- Third, are current counseling-testing program activities the best use of public health resources or are additional outreach strategies needed to target persons at highest risk of infection?

As these programs evolved in Minnesota over this 48-month observation period, they progressively served fewer clients at increased risk for HIV-1 infection and increasingly served clients at low risk for HIV-1 infection. A significant decrease in the number of homosexual/bisexual males seeking counseling and testing occurred over time. The HIV-1 antibody seroprevalence in homosexual/bisexual males seeking program services later in the observation period was lower than those seen early in the observation period, suggesting that men seen later were at lower risk for HIV-1 infection. In addition, 27 percent of the clients seen during this time period had been tested previously. Of these, most (97 percent) who initially tested negative were still negative for HIV-1 antibody on repeat testing. Some were clients who were instructed to return for follow-up testing to confirm their seronegative status, since they were in a potential "window-period" for HIV-1 antibody development based on their exposure history. Some of these repeat-test clients possibly were persons who continued to engage in high-risk behaviors. However, based on seroprevalence, it is likely that the majority of these repeat-test clients represent a low risk "worried well" population. Also, of the repeat-test clients who tested HIV-1 antibody positive, 175 (44 percent) had tested positive on a previous occasion, indicating that they likely were returning only to verify their infection status.

Counseling and testing site program policies remained stable during this time period. HIV-1 infection is a reportable condition in Minnesota and has been since shortly after the opening of these programs (October 1985), but clients are not denied services if they refuse to give a name or show proof of identity. This decrease in the number of persons at increased risk and corresponding increase in the number of persons at low risk for HIV-1 infection cannot be attributed to any change in program policy over the observation period. Also, similar trends demonstrating a decrease in HIV-1 seroprevalence at programs have been noted elsewhere in the country in states where HIV-1 infection is not reportable.[7–9]

Because long-term behavioral data were not collected at the sites, we have no direct measures of the effectiveness of our programs in promoting behavior change which prevents HIV-1 infection. Data on repeat-test clients can be used as an indirect measure. Of the 401 clients who were HIV-1 antibody positive on repeat testing, 107 (27 percent) had tested negative previously; data are not available on the time of the previous test for 28 (26 percent) of the clients. Eighteen (17 percent) had tested HIV-1 antibody negative within the previous six months and thus may represent persons who were infected but had not yet seroconverted at the time of the prior test. However, 61 (57 percent) of the 107 clients reported testing HIV-1 antibody negative in the previous seven months or longer and thus may represent new infections due to ongoing high-risk activity, despite having received risk reduction counseling at the time of previous testing.

Current data regarding knowledge of HIV-1 infection status and subsequent behavior change are unclear and suggest different effects. Knowledge of HIV-1 infection, particularly of a positive status, has been associated with a reduction of high-risk behavior in some studies[10–17]; however, only one of these studies examined persons tested at counseling and testing sites. Other studies have not demonstrated behavior changes associated with knowledge of HIV-1 infection.[18–21] These studies may not necessarily be extrapolated to persons tested at counseling and testing sites nor are there data regarding long-term behavior change after a single instance of counseling and testing.

In recent years, large amounts of federal and state dollars have gone into expanding and maintaining HIV-1 counseling and testing sites. Approximately $1.4 million was spent in Minnesota on this program at these two sites during this 48-month observation period. Despite expenditure of these resources, the trends identified here indicate that the program may not be reaching persons at highest risk and that providing free services to anyone requesting them may not be the best use of available public health dollars for AIDS prevention. In areas such as Minnesota where heterosexuals (other than intravenous drug users) are at low risk for HIV-1 infection, services need to be targeted to those at highest risk. In addition, counseling and testing sites might attract more persons at increased risk for HIV-1 infection by being linked to additional services such as HIV-1 disease diagnostic and treatment services.

Strategies to both identify and develop outreach programs for those at highest risk are also needed. Clinic-based sites likely to provide services to persons who are engaging in high-risk behaviors, such as chemical dependency treatment programs and clinics for treatment of sexually transmitted diseases, need to have well-developed risk assessment programs and on-site availability of HIV-1 antibody testing. Other outreach strategies, such as programs involving street workers, need to be expanded. Third-party partner notification programs, which actively and selectively reach those at highest risk for HIV-1 infection, need to be considered by state health departments as another outreach mechanism.[22–24] The data presented here may not be applicable to all areas of the country, particularly where HIV-1 infection has become endemic in certain racial and high-risk groups. Criteria can be developed locally to determine which persons are at highest risk for HIV-1 infection and testing sites can then be tailored accordingly.

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Potential for Laboratory Exposures to Biohazardous Agents Found in Blood

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Abstract: The magnitude of risk for occupational exposures to biohazardous agents found in blood was assessed by 800 environmental samples taken from a total of 10 clinical and research laboratories at the National Institutes of Health (NIH). Thirty-one samples from 11 work stations in three laboratories contained hepatitis B virus surface antigen (HBsAg). Observations of workers indicated that environmental contamination arose from several sources. Among the 11 work stations with HBsAg environmental samples, eight had high work loads, seven had inappropriate behaviors, and nine had flawed laboratory techniques. This information suggests that a multifactorial approach is needed to minimize the risk of laboratory-associated infections. (*Am J Public Health* 1990;80:423–427.)

Introduction

Hepatitis B virus (HBV) infection is among the most commonly reported laboratory-associated infections.[1–4] Laboratory workers in an urban medical center are at almost three times the risk of acquiring HBV infection than other hospital employees, based solely on exposure to patients' blood.[5] HBV infection is seven to 10 times more common among occupational groups in laboratory medicine than in the general public.[6,7] Workers' anxiety is heightened because of the concomitant risk of acquiring other infections with human immunodeficiency virus (HIV) and agent(s) of non-A, non-B hepatitis.

Strict adherence to biosafety guidelines minimizes the risk of acquiring laboratory-associated infections.[8,9] Immunization of workers is clearly the most effective means to control infection.[10] However, many employees are reluctant to be immunized and, as is the case with HIV, licensed vaccines are not always available. Thus, these types of laboratory-associated infections continue to occur,[11] despite the fact that most are preventable. Recently, universal blood and body-fluid precautions were recommended to further enhance workers' safety.[12]

This study dealt with an assessment of similarities and differences in clinical and research laboratory activities concurrent with a survey to detect hepatitis B surface antigen (HBsAg) in the environment. For this study, HBsAg served as a marker for contamination with pathogens found in human blood. Based on data from viral survivability on fomites, the selection of HBsAg as a marker for blood-borne contamination exaggerates the viability of a labile microorganism such as HBV. Nevertheless, contamination with HBsAg suggests a common source and depicts a potential for transmission of infection in the laboratory. Given this caveat, the purpose of this study was to elucidate factors associated with HBsAg in the environment and, therefore, potential exposures to blood-borne pathogens in biomedical laboratories at the authors' institution.

Methods

Recruitment and Stratification of Laboratories

Ten laboratories, representing a cross section of operations involving the use of human blood at the National Institutes of Health (NIH), were entered into a survey conducted from August 1984 through February 1985. We invited 119 laboratory workers to participate. Each volunteer was assigned a code to assure anonymity.

The laboratories surveyed fulfilled criteria for assignment to Biosafety Level II[8] and were chosen to represent a spectrum from high to low risk of workers handling blood containing HBV. The 10 participating laboratories were divided into 44 distinct work stations for purposes of data collection. Each of these stations was assigned a level of risk prior to environmental sampling based on the volume of blood processed and the hypothesis that the probability of detection would be greater in laboratories which processed blood drawn from individuals known to be infected with HBV than from those which excluded such blood samples. High-risk work stations were arbitrarily defined as those locations where the estimated volume of blood processed per week equaled or exceeded 1.5 L or the specimens received were from individuals known to belong to groups with a high prevalence of HBV infection. Stations assigned to the low-risk category were arbitrarily defined as areas where workers processed less than 1.5 L blood per week, with all blood drawn from donors known to be negative for HBsAg.

There were 75 hours of direct observation with the amount of time per laboratory depending on the continuity of tasks involving the use of blood. Observations were obtrusive because this approach allowed for interviews with workers while laboratory activities were assessed.[13] Data obtained in this manner may have been biased because of the worker's apprehension and/or attempts to modify unsafe practices. To minimize worker's reaction to being observed, the investigator sought to establish a good rapport with participants, reiterate key issues of the study, and assure the worker's anonymity. Stratification of risk in the work stations of the laboratories surveyed and the observation periods completed are listed on Table 1.

Laboratories 2 and 3 (Clinical Pathology and Transfusion Medicine Departments, respectively) provided diagnostic services for the Warren G. Magnuson Clinical Center (the hospital at NIH); the majority of work stations in these two laboratories were classified in the high-risk category. The work completed in Laboratory 4 was classified in the high-risk category because numerous specimens were analyzed for large-scale serological surveillance studies. Individuals in

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TABLE 1—Distribution of Laboratories, Type of Operations, Level of Work Station Risk, and Observations

Laboratory		Categories of Risk		Observations (h)[b]
Code	Operation	High (No.)[a]	Low (No.)[a]	
1	Immunological profiles	0	2	2
2	Diagnostic	14	1	28
3	Diagnostic	11	1	15
4	Immunoassays	2	0	4
5	Immunological profiles	3	0	5
6	Immunological profiles	0	1	2
7	Flow cytometry	1	0	1
8	Liver disease research	3	0	4
9	Tumor immunology	3	0	9
10	Virology research	2	0	5
Total		39	5	75

[a] Number of work stations.
[b] Hours.

Laboratories 5, 7, 8, and 10 processed blood specimens that were obtained from patients with chronic liver disease or HIV infection. In Laboratory 9, a tumor immunology unit, workers analyzed specimens from patients likely to have a history of multiple blood transfusions.

Assessment of Laboratory Operations

Methods for assessing skilled performance (both behaviors and techniques) were adapted from a task strategies approach.[14] To control for spurious conclusions due to inconsistencies with data collection,[13,14] all data were recorded by a single investigator who reviewed the study parameters frequently to control for observer drift. No intervention was attempted during the study.

Observation of activities involving the use of human blood were assessed by two indices. First, the frequencies of 13 different manipulations were each tallied to estimate the extent of workers handling blood that was required in order to complete a task. A task was defined as the cumulative manipulations generating a single result. The second index was the work load (the number of blood specimens processed per week). These indices allowed for comparison of laboratories performing quite different operations.

Behaviors were also assessed in an attempt to identify practices associated with an increased risk of contamination. Specific behaviors were assessed when a worker performed a movement either accidentally or intentionally that involved contact between human blood and the employee's skin, conjunctiva, or mouth. Each behavior was classified either as appropriate, if the practice was consistent with biosafety recommendations,[8,9] or inappropriate, if the worker was placed at risk of an exposure (e.g., mouth pipetting fluids or skin contact with blood) as a result of a breach in good laboratory practices. In the event of an overt accident the worker's behavior was judged to be appropriate only if the contamination occurred despite good laboratory practices (e.g., mechanical failure or poor instrument design). Observations were tallied prior to knowledge of sample testing.

Environmental Survey

Samples collected for the detection of HBsAg were processed in a manner consistent with Biosafety Level II guidelines.[8] HBsAg was eluted from various surfaces using a modification of a procedure described by Favero, *et al.*[15,16] The elution fluid (FBS/PBS) was prepared using 20 percent fetal bovine serum (Lot #14-501B: M. A. Bioproducts, Walkersville, MD) (FBS), phosphate-buffered saline (0.15 M NaCl and phosphate: M. A. Bioproducts, Walkersville, MD) (pH 7.4 ± 0.5) (PBS), and sodium azide (1.0 mg/ml) (Sigma Chemical, St. Louis, MO).

Samples were collected from work surfaces measuring approximately 35 cm^2 using moistened swabs and from swatches of gloves worn by workers or absorbent paper placed on the work surface prior to handling blood. Pledgets (each with a surface area measuring 5.0 cm × 7.5 cm) cut from Grade 304 stainless steel by the Biomedical Engineering and Instrumentation Branch, NIH, Bethesda, MD, were used for indirect sampling. These pledgets were placed in the laboratories either in the working area or, for negative controls, outside the working space. Negative controls were graded unsatisfactory if the pledgets were either moved or visibly soiled during the exposure period. Metal pledgets were used because preliminary modeling studies demonstrated that nanogram amounts of HBsAg could be eluted from four surfaces (absorbent paper, glass, polystyrene, and stainless steel) with recovery rates averaging 20 percent, 77 percent, 82 percent, and 96 percent, respectively. Other data have shown that HBsAg can be recovered from metal surfaces with greater than 99 percent efficiency.[15] All samples were eluted in 1.5 ml FBS/PBS and stored at 4°C.

Hepatitis B Surface Antigen Detection

Samples were screened for the presence of HBsAg using a radioimmunoassay (Ausria II: Abbott Laboratories, Chicago, IL) according to the manufacturer's instructions with these modifications: a) each control and sample was assayed in duplicate using 200 ul eluate per well; b) six distinct negative environmental controls were assayed with each run; and c) each assay tube was counted for 10 min. Tubes were counted in a gamma spectrophotometer (Tracor Analytic, Model 1185: Nuclear-Chicago, Des Plains, IL) with a calculated counting efficiency for I-125 of 83 percent. The average cpm for each paired assay was calculated. Samples with discrepant paired results were retested. Those samples with repeatedly discrepant results were excluded from the study.

A sample was classified as reactive if the cpm for the sample was greater than three standard deviations from the mean cpm for the negative controls. The coefficient of variation (CV) was calculated and the mean of the negative controls for each screening run (values ranged from 3 percent to 17 percent with a mean of 7 percent).

Reactive samples were confirmed using specific neutralizing antibody for HBsAg (anti-HBsAg) (Ausria II Confirmatory Neutralizing Kit: Abbott Laboratories, Chicago, IL). Appendix A outlines the method used for the confirmation assay. A sample was confirmed HBsAg positive if the reduction in the cpm of the sample reacted with anti-HBsAg was greater than 50 percent compared with the sample reacted with FBS/PBS and if these latter cpm were between the range of 50 percent to 150 percent of the mean cpm of the negative control. Reactive samples with cpm less than 1.5 times the mean cpm of the negative control were diluted beyond the sensitivity of this latter assay and were interpreted as borderline. Within run variation for the confirmation runs was expressed as the CV calculated for the negative control (values ranged from 39 percent to 59 percent with a mean of 4 percent).

Statistics

The association of the potential factors with HBsAg was assessed by determining odds ratios and 95 percent confidence intervals.[17]

Results

The results for the HBsAg sampling data are summarized on Table 2. These data show that our hypothesis was correct: high-risk operations do differ from low-risk operations in terms of the potential for blood-borne contamination in the laboratory. All environmental samples collected from the low-risk work stations were negative for HBsAg. A total of 69 samples collected from high-risk work stations in Laboratories 2–4 and 8–10 were reactive for HBsAg on the initial screening. Except for the tumor immunology research unit, these laboratories routinely processed blood drawn from individuals chronically infected with blood-borne pathogens. Approximately 55 percent (38/69) of the reactive samples were graded as borderline (low cpm for positive samples overlapped with high cpm for negative controls). Insufficient eluate volumes prohibited further analysis of these borderline samples. Thus, we analyzed a total of 800 samples, of which 31 (3.9 percent) were HBsAg positive, 38 (4.8 percent) were borderline, and 731 (91.4 percent) were negative.

Samples confirmed positive for HBsAg were collected from Laboratories 2, 3, and 8. The HBsAg positivity from these laboratories ranged from a high of 17.8 percent (8/45) (liver disease research unit) to a low of 1.4 percent (2/139) (Transfusion Medicine). The frequency of obtaining positive samples in Clinical Pathology was 8.6 percent (21/245).

Factors Associated with HBsAg in High-Risk Work Stations

Individuals who worked in nine different locations in the Clinical Chemistry and Hematology Services had the highest work loads (average 1,430 specimens/week: range from 210–2,700). Workers in the Transfusion Medicine Department averaged 106 specimens per week. The liver disease unit, the laboratory with the greatest percent of samples positive for HBsAg, averaged 27 specimens per week. Work loads for Laboratories 4 and 7 averaged 1500 and 350, respectively. The remaining laboratories analyzed fewer than 60 specimens per week (range from 13–60).

Inappropriate behaviors associated with an obvious risk for cutaneous contact with blood were identified at nearly 40 percent (15/39) of the areas surveyed. One example of these inappropriate work practices which occurred was that workers routinely handled contaminated fomites without wearing gloves. Despite the presence of warning labels (e.g., blood and body-fluid precautions) on some specimens, workers at 60 percent (9/15) of these same areas were not motivated to wear gloves. Workers who consistently wore gloves when working with blood (whether a warning label was present or absent) were observed at only 33 percent (13/39) of the stations. Surprisingly, two workers who routinely analyzed blood donned gloves only when they handled blood with a warning label.

TABLE 2—Summary of HBsAg Environmental Sampling Data

Laboratory No.	No. Tested[a] %[b]	HBsAg Reactive			
		Initial Screen		Confirmed	
		No.	%[b]	No.	%[b]
1	87	0	0.0	0	0.0
2	245	44	18.0	21	8.6
3	139	11	7.9	2	1.4
4	46	1	2.2	0	0.0
5	90	0	0.0	0	0.0
6	19	0	0.0	0	0.0
7	14	0	0.0	0	0.0
8	45	10	22.2	8	17.8
9	47	1	2.1	0	0.0
10	68	2	2.9	0	0.0
Total	800	69	8.6	31	3.9

[a] No. = Number.
[b] Percent of total tested.

Four workers committed flagrant violations of good work practices during the observations. These individuals either mouth-pipetted serum or live virus suspension or splashed milliliter amounts of body-fluid onto their skin. They were also observed to disregard basic hygienic practices (e.g., placing laboratory marking pens in the mouth) and misused equipment. Examples of other unsafe work practices observed in this study included: sonicating infectious materials without physical containment and hand braking a centrifuge rotor. Curiously, three employees acknowledged that they were aware of the risks associated with their actions.

Table 3 shows the frequencies and crude odds ratios of HBsAg contamination in 39 high-risk work stations. High work load and flawed technique were associated with the greatest risk of contamination, whereas there was a borderline association with inappropriate behavior (Table 3). Flawed techniques were of two kinds: those related to instrument design and those related to procedures which were intrinsically defective (i.e., workers may have been following the procedure as described, yet blood spills routinely occurred).

Approximately 85 instruments that workers used during the study were sampled for HBsAg. The design of these instruments was considered flawed if a mechanical feature generated a spatter of blood or necessitated frequent maintenance of components contaminated with body-fluids. HBsAg was detected on surfaces near several instruments where blood was spattered: Sequential Multiple Analyzer and Computer, Model SMAC II Systems: Technicon Instruments Corporation, Tarrytown, NY; Automated pH/Blood Gas System, Model 175: Corning Medical and Scientific, Medfield, MA: Pentawash II: Abbott Laboratories, Chicago, IL; and Coulter Counter, Model S-Plus System, Coulter Electronics, Inc., Hialeah, FL. Similarly, HBsAg was detected on instruments designed with samplers that required external washing. Compared with all other types of instruments surveyed in our laboratories, the continuous flow analyzers were the most vulnerable to malfunctions of

TABLE 3—Comparison of Factors with HBV Contamination in Laboratory Work Stations Assigned to the High-Risk Category

Factor	HBsAg Survey	Odds Ratio and 95% Confidence Interval
	Number of Stations with HBsAg/Total	
Work load (sera/wk)		
≤150	3/24	
>150	8/15	8.00 (1.55, 31.19)
Behavior		
Appropriate	4/24	
Inappropriate	7/15	4.38 (0.97, 16.61)
Containment Measures		
Appropriate Safeguards	2/22	
Flawed Techniques	9/17	11.25 (1.84, 46.06)
Complexity (manipulations/task)		
<30	10/29	
≥30	1/10	0.21 (0.04, 1.92)

intricate serum-containing components. All of these design features were associated with an increased occurrence of contamination in our laboratories.

Instruments with discrete reaction containers and samplers with fluid sensors that do not require external wash steps minimize contamination. Visible soiling was not observed and HBsAg surveys were negative after operating the following instruments: Automatic Clinical Analyzer III: Dupont Instruments, Wilmington, DE; Automated STAT/Routine Analyzer-8: Beckman Instruments, Brea, CA: Immunochemistry Systems Analyzer II: Beckman Instruments, Brea, CA; and Cobas-BIO: Roche Analytic Instruments, Inc., Nutley, NJ.

Techniques which frequently resulted in blood spills, despite observations that workers adhered to written procedures, were considered defective because these methods did not reasonably ensure containment of biohazardous fluids. They were frequently associated with a need for repetitious separations and intermittent mixing of blood without adequate measures to contain spills. Both image and flow cytometric systems were identified as techniques with the most frequent occurrence of overt spills. These types of systems were associated with the presence of HBsAg in Laboratories 2, 3, and 8.

To approximate the risk of contamination when processing blood contained in vacuum-sealed tubes, samples were collected from a device designed for removing rubber stoppers from tubes. Operation of the device entails aligning the tube into an opening and, then, mechanically directing a metal arm to swing against the stopper forcing it to pop off simulating the manual technique. HBsAg was detected repeatedly on the rim of the device opening where spatter occurred. These data suggest that removing stoppers by either method is a subtle source of contamination. We endorse the technique of "wrap and snap" to minimize spatter of blood.

Tasks with few manipulations usually involved processing small volumes of blood using microscale techniques (both manual and automated). Automated instruments were used in 8 of 15 stations surveyed in Clinical Pathology compared with only 2 of 11 stations in Transfusion Medicine. Exclusive of the flow cytometry service, only manual techniques (e.g., immune profiles) were observed in the remaining research laboratories. Automated assays using less than 15 ml of blood averaged three manipulations per task: whereas, comparable manual assays averaged 10 manipulations per task. Tasks using greater than 1 L of blood required the most handling (average 30 manipulations/task: range from 6–92). These latter procedures were typically performed in research laboratories.

To examine the association between different tasks and contamination, we arbitrarily grouped each work station into those involving less than 30 manipulations and another group with 30 and more activities (complex tasks). In contrast to the factors mentioned earlier, complexity appeared to be inversely associated with contamination (OR = 0.21, 95% CI = 0.04, 1.92) (Table 3). A review of the data revealed that complexity was inversely correlated with work load (no complex tasks were performed in stations with high work load).

To further examine the predictors of blood-borne contamination, logistic regression analysis was performed to determine the effects of the factors shown on Table 3 while controlling for the effects of each other. Because of the negative correlation between complexity and high work load, we chose to remove complexity from further analysis. In the analysis with the three remaining factors, the point estimates for the adjusted odds ratio for flawed technique decreased to 9.78 (95 percent CI = 1.46, 65.49) and was the factor most strongly associated with HBsAg. Only a modest decrease in the adjusted odds ratios was observed for high work load (OR = 5.06, 95% CI = 0.80, 31.96) and inappropriate behavior (OR = 2.75, 95% CI = 0.44, 17.4) compared with the crude odds ratios. In a similar analysis with only two variables, both high work load (OR = 6.82, 95 percent CI = 1.73, 39.66) and flawed technique (OR = 9.78, 95 percent CI = 1.52, 63.01) were associated with an increased risk of blood-borne contamination. (Data available on request to authors.)

Discussion

Data from our study are consistent with the concept that workers who process blood from numerous donors (high work loads typically associated with diagnostic testing) increase their risk of occupational exposures to blood-borne pathogens. However, a high prevalence of HBV infection among blood donors can also contribute to risk. For example, the liver disease research unit had the highest HBsAg positivity rate (17.8 percent) but only averaged 27 specimens/week. One-third of the specimens submitted to this laboratory were from individuals known to be chronic carriers of HBV, compared with the other laboratories surveyed where fewer than one-tenth of the specimens were obtained from infected patients. It is important to note, however, that an analogous risk can be identified in any laboratory where serial specimens from infected patients are analyzed (e.g., therapeutic drug monitoring, diagnostic microbiology, and similar types of research involving HBV or HIV). In support of this observation is the recommendation that the function or nature of the laboratory should also be considered in the hazard assessment to protect these workers.[8]

Inappropriate behaviors (a second factor associated with HBsAg) were noted when workers disregarded standard practices or misused equipment. Somewhat alarming were observations that individuals at almost 50 percent of the high-risk areas experienced a seemingly innocuous, but preventable cutaneous contact with blood as a result of poor work practices. Complacency regarding the wearing of gloves while handling contaminated fomites was cited as the most frequent cause of these exposures.

During 75 hours of observation, only four overt accidents resulting in contact between the employees' protective garb and blood were observed. No accidental parenteral inoculations occurred and none of the four accidents was associated with inappropriate behaviors. These data can be contrasted with our observations cited above that poor work practices were associated with the employee experiencing cutaneous contact with blood. Based on these data, we concluded that the majority of exposures to blood that occur in the laboratory are occult in nature, as well as, preventable. Further, the risk of processing blood containing a pathogen is likely to be underestimated.[18] Despite the presence of warning labels on specimens submitted to NIH laboratories, during our observations, the majority of workers who engaged in poor work practices were not motivated by these labels to modify their inappropriate behaviors. Equally important was our observation that none of the individuals with good work practices experienced an adverse exposure to blood. In combination, these data reinforce the concept that workers who handle human blood strictly adhere to precautions as an effective means to reduce their risk of laboratory-associated infection. These practices were observed prior to the implementation of universal precautions in our hospital.[12]

Our results indicate that automation contributed to contamination only when the design of certain instrument components was flawed with respect to containment.[19] Just as good engineering design can be negated by poor work practices, flawed design (the factor associated with the greatest risk of contamination) can also overcome good work practices.[9] Our study demonstrates the need to modify instrument components which represent an increased risk for contamination and, perhaps, infection among laboratory workers.

Manual techniques such as cytometric systems that involved sequential manipulations of specimens represented an increased risk for contamination in our laboratories. The frequent presence of HBsAg in these areas indicates that blood-borne contamination is predictable for many of these procedures. We concur that workers should modify these techniques in order to enhance their safety.[9]

In our study, the complexity of laboratory tasks (in terms of cumulative manipulations per task) was weakly associated with the risk for HBsAg. Also, we observed a negative correlation between complexity and high work load which probably reflected differences in the processing of specimens for research protocols compared to diagnostic testing. We submit, therefore, that the complexity of tasks performed represents minimal risk to the worker and recommend that predictors associated with greater risk of blood-borne contamination be given priority in the assessment of laboratory safety. Use of these data and previously published data[2,8,12] in designing interventions to eliminate adverse exposures to blood may prevent implementing strategies that unnecessarily disrupt the laboratory while not enhancing worker safety.

APPENDIX A

Method Used for HBsAg Confirmation Assay

Modifications to the manufacturer's instructions (Ausria II Confirmatory Neutralizing Kit: Abbott Laboratories, Chicago, IL) were as follows. The HBsAg positive control (Ausria II: Abbott Laboratories, Chicago, IL) was processed in the same manner as the samples. After the addition of a polystyrene bead coated with guinea pig anti-HBsAg, 150 µl of eluate or HBsAg positive control were dispensed to each of four wells. The negative control (FBS/PBS) was assayed in replicates of six using 150 µl per well. After the initial wash step, 150 µl of FBS/PBS (diluted 1:5 with 0.9% NaCl) were added to the negative control wells. Each tray was incubated (25°C for 16–24 h). Each well was then washed four times using 5 ml distilled water per rinse. A 150 µl volume of confirmatory anti-HBsAg (diluted 1:5 with 0.9% NaCl) were added to two of the four wells and 150 µl of FBS/PBS (diluted 1:5 with 0.9% NaCl) were added to the remaining two wells of each set of assays and to each of the negative control wells. Each tray was incubated (45°C for 1 h). Without aspirating the contents of the well, 200 µl of I-125 anti-HBsAg were then added to all of the wells and allowed to react (45°C for 1 h). After repeating the wash step, each bead was counted.

The average cpm was calculated for each replicate assay that reacted with either anti-HBsAg or FBS/PBS. The percent reduction in the cpm for the positive control and each sample was calculated using the following formula:

$$\% \text{ Reduction} = \frac{\{\text{cpm}(\text{sample} + \text{FBS/PBS}) - \text{cpm}(\text{sample} + \text{anti-HBsAg})\} \times 100}{\{\text{cpm}(\text{sample} + \text{FBS/PBS}) - \text{cpm}(\text{negative control})\}}$$

Validity of a confirmation run was based on two arbitrary criteria: 1) neutralization with anti-HBsAg had to result in a greater than 50 percent reduction in the cpm for the positive control when compared to the cpm of the control reacted with I-125 anti-HBsAg (the parameter was >97 percent for each run); and 2) no more than two replicates of the negative control could be considered aberrant. This latter event did not occur. However a single value was greater than 1.5 times the mean of the negative control for three of the five runs. In each case, the aberrant value was excluded and the mean recalculated.

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Recovery of Hepatitis A Virus from a Water Supply Responsible for a Common Source Outbreak of Hepatitis A

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Harold S. Margolis, MD, Howard A. Fields, PhD, Thomas W. McKinley, MPH,
Charles P. Gerba, PhD, James E. Maynard, MD, PhD, and R. Keith Sikes, DVM, MPH

Abstract: An outbreak of hepatitis A occurred in a north Georgia trailer park served by a private well. Of 18 residents who were serosusceptible to hepatitis A virus (HAV), 16 (89%) developed hepatitis A. Well water samples were collected 3 months after illness onset in the index case and 28 days after illness onset in the last trailer park resident. Hepatitis A virus antigen (HAVAg) was detected in the samples by enzyme immunoassay from three of the five cell lines following two 30-day passages and from a fourth cell line following a third passage of 21 days. (*Am J Public Health* 1990; 80:428–430.)

Introduction

Infection with hepatitis A virus (HAV) accounted for 39 percent of the 59,633 cases of acute viral hepatitis reported in the United States in 1985.[1] Virus transmission occurs primarily through the fecal-oral route, resulting in both sporadic cases and common source outbreaks.[2] Common source outbreaks of hepatitis A have been associated with ingestion of contaminated food[3,4] or consumption of inadequately treated drinking water.[5–9] The incrimination of food and water in HAV transmission has largely been based on the epidemiologic characteristics of the outbreak and serologic evidence that the resulting viral hepatitis was type A. Only recently has HAV been recovered from water associated with hepatitis A outbreaks.[10]

HAV is classified as a picornavirus, and most wild-type HAV does not produce a cytopathic effect, although infectious units can be quantitated by the radioimmunofocus assay (RIFA).[11] We describe the cell culture isolation of HAV from drinking water implicated as the source of virus transmission in a common source outbreak.

Methods

Epidemiologic Investigation

A 21-year-old man had onset of serologically confirmed hepatitis A on July 10, 1982 while living from July 9 through July 23 in a trailer park in Bartow County, Georgia. Approximately three to nine weeks following onset of illness, other residents of the trailer park developed hepatitis A. All cases of hepatitis A during 1982 were identified using county health department case reports and laboratory records for hepatitis A serologic testing. Each trailer park resident was interviewed and a blood sample obtained. All serum samples were tested by radioimmunoassay for total and IgM-specific antibody to HAV (HAVAB and HAVAB-M, Abbott Laboratories,* North Chicago, IL).

Case Definition

A case was defined by a positive serologic test for IgM anti-HAV. One trailer park resident who declined a blood test but reported a history of jaundice and abdominal pain was also included.

Water Sampling

Tap water from the trailers was tested for fecal coliform bacteria by standard methods.[12] Fluorescein dye was discharged from the toilet of the index case, and tap water samples were collected after 19 and 81 days.

Three months after onset of illness in the index case, 400 and 1600 liters of ground water from the well pump supplying the park were passed through an electropositive charge-modified, 25-cm long, cellulose-diatomaceous earth composite, microporous filter (Virosorb 1MDS; Cuno, Meriden, CT).

Cell Cultures

Adsorbed viruses were eluted from the filter cartridges with 1 liter of 3 percent beef extract adjusted to pH 10, and after elution, sufficient 1M glycine was added to lower the pH to 3.5. The flocculated organic material was processed according to the method of Katzenelson, *et al.*[13] Virus culture was carried out in four monkey kidney cell lines: BGM, BS-C-1, FRhK-4 (American Type Culture Collection, Rockville, MD) and MA104 (Microbiological Associates, Bethesda, MD); and in primary Rhesus monkey kidney (PMK) cells.

The concentrated sample was filter sterilized directly onto the cell monolayer, and virus was adsorbed at 37°C for two hours with periodic rocking. PBS mock-infected cell cultures were included for controls. All cells except FRhK-4 were grown in Eagle's minimum essential medium containing Earle's balanced salts solution and supplemented with 10 percent fetal bovine serum, 25 mM HEPES (N-2-hydroxyethylpiperazine-N'-2-ethanesulfonic acid), 2 mM L-glutamine, 0.1 mM nonessential amino acids, 0.2 percent sodium bicarbonate, and antibiotics (20 mcg/ml gentamycin, 100 mcg/ml streptomycin, and 100 U/ml penicillin). FRhK-4 cells were grown in William's medium E as previously described.[14] Maintenance media consisted of the same basal medium

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*Use of trade names is for identification purposes only and does not imply endorsement by the Public Health Service, US Department of Health and Human Services, or the Georgia Department of Human Resources.

containing 2% fetal bovine serum. All media and media components were obtained from Gibco Laboratories (Grand Island, NY).

Continuous cell lines were passaged twice at 30-day intervals and four times at 21-day intervals. Virus was recovered from PMK cells by trypsinizing, freeze-thawing, sonication, extraction with 2,3,5 trichloro-trifluoroethane (Dupont Chemicals, Wilmington, DE), and passaged acutely.

Assays for HAV

Trypsinized cell samples were extracted as described above and assayed for HAVAg with a double-antibody sandwich enzyme immunoassay (EIA), as previously described.[15] A ratio of positive to negative values (P/N) >2.1 was considered positive for HAVAg.

Infectious units of HAV were quantitated by the RIFA, using both human and rabbit anti-HAV.[11] HAV titers were expressed as radioimmunofocus-units (RFU) per ml.

Results

Epidemiologic Investigation

Thirty-five cases of hepatitis A with onset of illness from July 10 to November 24, 1982 were identified in the county. Before July 10, there had been no reported hepatitis A cases in the county during 1982. Of the 35 cases, 17 were residents of the trailer park. In 31 cases the date of onset of symptoms could be recalled (Figure 1). Among the 13 residents of the trailer park who could recall date of onset of illness, 12 became ill within 22 to 41 days following disease onset in the index case.

The index case was a 21-year-old male who lived with a family in the trailer park from July 9 to July 23. He denied contact with the seven other families in the park. Of the 30 trailer park residents, seven had serologic evidence of prior exposure to hepatitis A. The five persons living with the index case received immunoglobulin prophylaxis, and one became IgM anti-HAV positive. Among the 18 remaining trailer park residents considered serosusceptible to hepatitis A, 16 developed hepatitis A (89 percent attack rate). The epidemiologic investigation revealed no common social activity, meal, or food which could account for the outbreak. The only two serosusceptible individuals who did not become ill were 11 and 18 months of age, and their mothers were positive for total anti-HAV but negative for IgM anti-HAV.

Of the 18 cases who were not residents of the trailer park (including the index case), 11 had visited the trailer park and drunk water during their visit (Figure 1). Another four patients had contact with cases only outside the trailer park, while three had no known contact with other cases.

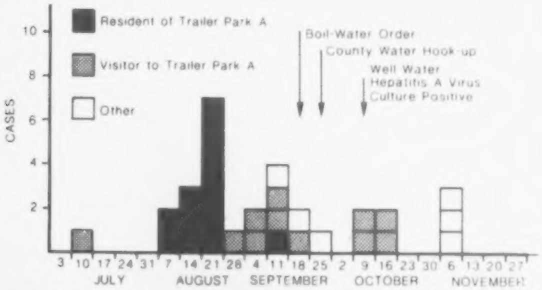

FIGURE 1—Onset of Reported Hepatitis A Cases with Known Date of Onset in Bartow County, Georgia, July through November 1982. (A county order to boil water was issued September 17. The trailer park was connected to the county drinking water supply on September 24. The hepatitis A virus-containing water samples were collected from the implicated well on October 6.)

Environmental Investigation

The trailer park's sole water supply was a 15-cm diameter drilled well having a depth of 24 to 30 m, encased within steel pipe. The trailers encircled the well at a distance of 18 to 60 m, and water was distributed through a 2-cm galvanized pipe buried approximately 50 cm below the ground surface. Wastewater from each trailer was treated by individual septic tank systems having gravity feed, subsurface drainage tile distribution networks; each septic tank system was separated from the well by a distance of 30 to 60 m.

The concentration of fecal coliforms in tap water ranged from zero to greater than 16 per 100 ml and in water collected from the wellhead was 2 per 100 ml. Fluorescein dye was flushed from the toilet of the trailer in which the index case had lived, but could not be detected on days 19 and 81 after charging the system.

A boil-water order was issued on September 17, and on September 24 the well was disconnected, and the water distribution system was connected to the county water supply.

Virus Assays

HAVAg could not be detected in the ground water concentrates by EIA. Following amplification in cell culture, HAVAg was detected in the monkey kidney cell lines (Table 1). After primary cell adaptation to FRhK-4 and PMK cells, foci of HAV were detected by RIFA (Figure 2). Passage in FRhK-4 and PMK cells yielded titers of 3.0×10^7 RFU/ml and 2.0×10^7 RFU/ml, respectively.

Discussion

The importance of HAV transmission by water is probably underestimated since outbreaks appear to be infrequently reported, poorly documented, and the virus has been only rarely isolated from this source.[9,10,16] Using currently available methods for enterovirus detection, HAV can be concentrated 100- to 10,000-fold onto microporous filters and concentrated another 10- to 100-fold by organic flocculation.[10,16,17] In one study in which these concentration techniques were used, samples of ground water implicated in a large outbreak of hepatitis A revealed low levels of HAVAg by RIA.[16] In contrast to a previously reported outbreak where HAV was isolated from contaminated water,[10] the fecal coliform counts obtained at the peak of the presumed period of contamination in the present outbreak were 10- to 100-fold lower, suggesting

TABLE 1—Detection by Enzyme Immunoassay of Hepatitis A Virus Recovered from Ground Water following Growth in Monkey Kidney Cells. (Ground water samples were collected October 6, 1982 from a well in a trailer park in Bartow County, Georgia.)

Monkey Kidney Cell Line	Hepatitis A Virus Antigen by Enzyme Immunoassay[a] by Passage Level					
Passage:	1[b]	2	3[c]	4	5	6
MA104	<1	<1	<1	<1	<1	<1
BGM	<1	1.1	2.4	3.6	6.9	21.0
BS-C-1	<1	2.9	8.5	11.0	21.0	>115.0
PMK	1.5	3.2	8.5	11.0	41.3	>115.0
FRhK-4	<1	2.4	5.6	10.6	57.0	160.0

[a] HAVAg expressed as positive/negative (P/N) ratios in undiluted cell lysates
[b] Passages 1 and 2 at 30 days
[c] Passages 3–6 at 21 days

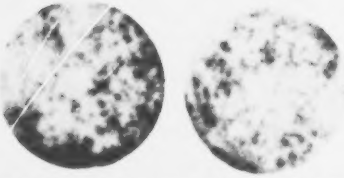

FIGURE 2—Foci Produced in BS-C-1 Cells in the RIFA from HAV Isolated from Concentrated Ground Water in Bartow County, Georgia following Initial Isolation in FRhK-4 Cells. (Panels from left to right represent increasing dilutions of the recovered HAV.)

that low levels of ground water contamination can produce significant outbreaks of hepatitis A.

This outbreak also demonstrated the stability of HAV in untreated water. The water samples from which the virus was isolated were obtained three months after the onset of symptoms in the index case and 28 days after the onset of symptoms in the last trailer park resident (Figure 1). The recovery of HAV suggests that the contaminating virus had continued access and persisted in the aquifer supplying the well although it cannot be ruled out that fourth generation cases (illness onsets from October 4 to 13) who were visitors to the trailer park, may have recontaminated the aquifer. In addition, the water concentrates were stored at 4°C for eight months before they were inoculated into cell culture.

While the incidence of HAV infection has been decreasing in the US over the past 10 years, it continues to be the second leading cause of viral hepatitis. Although the majority of the population utilizes municipal drinking water, it is evident that well water contamination with HAV continues to occur and that this source of infection is not widely appreciated.

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Review Criteria for Schools of Public Health Funding, Project Grants

Final review criteria, funding priority, and special consideration for special project grants to Schools of Public Health were published in the February 8 issue of the *Federal Register*. Among the review criteria to be considered is whether the project contains realistic objectives as described in *The Future of Public Health, Objectives for the Nation* and the *Sixth Report to the President and Congress on the Status of Health Personnel in the United States*.

The funding priority is designed to promote health initiatives that will encourage and increase involvement of underrepresented minority health trainees. Special consideration will be given to applications that address five issues:

- Substance Abuse
- HIV/AIDS
- Geriatrics
- Environmental health
- Training students to work in public health agencies

For further information, contact Blake Crawford, HRSA, US Public Health Service, DHHS, 5600 Fishers Lane, Room 14-43, Rockville, MD 20857; (301) 443-3376.

A Rural Outbreak of Legionnaires' Disease Linked to Visiting a Retail Store

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Abstract: Between May 7 and June 7, 1986, 27 residents of a rural county in Maryland developed legionellosis, and two died. *Legionella pneumophila* serogroup 1 was cultured from the sputum of two patients and identified in lung tissue of a third patient by direct fluorescent antibody staining. An additional 11 patients had four-fold rises in antibody titer to *L. pneumophila*, and 13 had single titers ≥1:256. To determine risk factors for disease, we performed a case-control study. Twelve of 16 case-patients reported visiting store A in the two weeks before onset of illness compared with four of 28 control-patients. A serologic survey of employees showed that employees of store A were 3.63 times more likely than control employees to have titers of antibody to *L. pneumophila* ≥1:256 (95% confidence intervals 0.8, 16.7). Cultures of soil specimens, samples of water from the hot water system of store A and from stagnant ponds near store A collected five weeks after the end of the outbreak were negative for *Legionella* species. Store A was adjacent to a site of excavation and construction during May 1986, when the community was experiencing an extended drought. This investigation suggests that exposure to excavation and construction activity may be a risk factor for legionellosis. *(Am J Public Health* 1990; 80:431–434.)

Introduction

Investigators agree that *Legionella pneumophila* is ubiquitous in nature, but they disagree about the importance of various modes of transmission.[1] Attention has focused on inhalation of organisms from water droplets in aerosols and aspiration of contaminated potable water as predominant modes of transmission. These conventional assumptions of how legionellae are transmitted are inadequate to explain outbreaks associated with construction and excavation when aerosols or potable water sources are not implicated. In this report we describe a community outbreak of legionellosis that was not associated with an aerosol or contaminated potable water. Exposures to excavation and construction were the only identifiable risk factors.

During the third week of May 1986, the staff radiologist of a community-hospital in a rural county of Maryland recognized an unusual cluster of cases of severe community-acquired pneumonia. The Maryland State Health Department and local county health department were notified and an investigation was initiated. Review of chest radiograph reports confirmed an increase in the number of both hospitalized patients and outpatients with radiographic evidence of pneumonia in May 1986 compared with April 1986, April 1985, and May 1985. Cases identified by this review had a wide range of clinical findings, and initial microbiologic studies did not implicate any single pathogen.

Methods

We defined cases of outbreak-associated Legionnaires' disease as follows: onset of symptoms consistent with pneumonia during May 1986, residence in county A, and either isolation of *L. pneumophila* serogroup 1 from a clinical specimen, fluorescent antibody staining of organisms from tissue, four-fold rise in antibody titer to *L. pneumophila* serogroup 1 to a titer ≥128, or a single antibody titer ≥256.

Cases were ascertained by reviewing all reports of chest radiographs performed on inpatients and outpatients during May and June 1986 at the only hospital in the county, clinical histories recorded in charts of patients reported to have pneumonia, and the reports by community physicians of patients with pneumonia. Convalescent-phase serum specimens were obtained on 52 suspect outbreak patients with pneumonia; for 41 patients an acute-phase serum specimen was also obtained.

Isolates of *L. pneumophila* obtained from patients were subtyped by the pattern of binding of a panel of monoclonal antibodies to *L. pneumophila*.[2]

To determine risk factors associated with illness, we performed a case-control study. Since laboratory data were unavailable at the time the case-control study was performed, we used a clinical case definition—temperature ≥101°F, white blood cell count ≥11,000 cells/mm^3, and dense pulmonary infiltrate by chest radiograph. A questionnaire was administered to each case-patient by phoned interview that lasted approximately 10 minutes. We asked whether the patient had visited any of eight specific locations in county A in the two weeks before illness onset and if so, on what date and at what time; whether the patient would have visited these locations during a usual two-week period; underlying medical problems; employment history during the two weeks before illness onset; and use of home air conditioning. At the end of the interview, case-patients were asked to identify two friends of the same sex as themselves and within 10 years of age to serve as control-subjects. Control-subjects were then telephoned, and the same questionnaire was administered. Risk factors were analyzed in a univariate fashion using the Pike-Morrow method for a variable number of controls per case without breaking the match.[3]

Results of the case-control study were used to focus subsequent investigation. Employees working within one-half mile of the retail store associated with cases of Legionnaires' disease, including the store itself, were asked to complete a questionnaire that requested information about recent illnesses and days missed from work. Serum speci-

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mens were obtained from employees working in the implicated area and from control employees of similar age working at a similar retail store located 15 miles from the implicated area who volunteered to be bled. Antibody titers to *L. pneumophila* serogroup 1 were determined by standard methods.[4] Titers were compared by Fishers' exact test.[5]

Water samples were obtained from potential reservoirs of *Legionella* and cultured according to established protocols.[6] Methods used included filter concentration, acid treatment, and use of antimicrobial containing media.Soil samples were also obtained, hydrated in sterile water, and cultured by the same methods.

Records of precipitation and wind direction collected 3.7 miles from the implicated area were obtained from the National Oceanographic and Atmospheric Administration (Asheville, North Carolina), and the environment of the area associated with acquisition of Legionnaires' disease was inspected to determine whether any aerosol generating devices could be identified.

Results

Twenty-seven residents of county A developed Legionnaires' disease during May 1986, and two died. The dates of the onset of symptoms ranged from May 11 to May 28 (Figure 1). *L. pneumophila* serogroup 1 was isolated from two patients' sputa; lung tissue from a third patient, obtained post mortem, gave a positive direct fluorescent antibody stain for *L. pneumophila* serogroup 1; 11 patients had four-fold rises in reciprocal antibody titer to *L. pneumophila* serogroup 1, and 13 had reciprocal titers ≥256 (four patients had reciprocal titers of 256). The two isolates of *L. pneumophila* from patients with Legionnaires' disease gave the same monoclonal antibody binding pattern, binding to monoclonal antibodies 1, 2, 5, 7. Patients with Legionnaires' disease ranged in age from 27 to 80 years (median 51). Age-specific attack rates of disease, using 1982 estimates of the county's population, showed that although many young adults were affected, attack rates tended to increase with age (Table 1). Only 17.6 percent of the cases were in women. The locations of patients' residences were spread over a large geographic area, virtually the entire county, without clustering of cases by census tract (Figure 2).

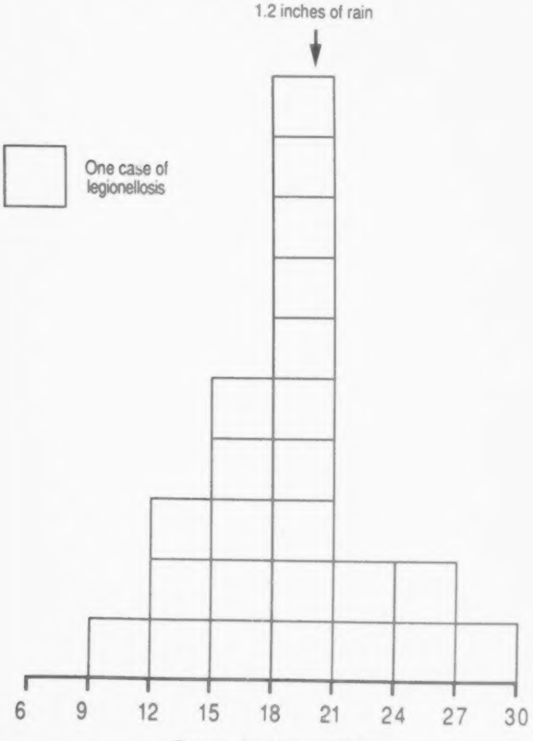

FIGURE 1—Date of Symptom Onset of Cases of Legionnaires' Disease (onset date unknown for 3 persons)

In the case-control study, 16 patients with Legionnaires' disease and 28 controls were interviewed. Of the 27 cases of Legionnaires' disease, 10 did not meet the clinical criteria required for enrollment in the case-control study, and one could not be located. Twelve cases and four controls had visited store A in the two weeks before illness onset (lower bound 95% confidence interval 3.42). No other potential risk factor was significantly more common for the cases.

Store A is a large single room retail store employing approximately 100 employees situated in a sparsely populated area of County A with few nearby residences but some recent commercial development. Several hundred customers visit store A per day. Approximately 100 yards south of Store A, excavation and construction of a shopping mall took place during the month of the epidemic. No cooling towers or wet-heat rejection sytems were present within one-half mile of store A.

The survey of Store A employees revealed that one of 90 had had pneumonia during May. This individual had a reciprocal antibody titer to *L. pneumophila* serogroup 1 of 256 five weeks following the onset of his illness. Greater prevalence of detectable antibody to *Legionella* and higher titers were seen in Store A employees than in persons working in store B and near Store A (Table 2). Store A employees were 3.63 times as likely to have reciprocal titers of 256 or more (95% CI = 0.8, 16.7). Because of the delay between the survey and the end of the outbreak, it was not possible to obtain serum specimens or interviews from persons working at the construction site during May. No *Legionella spp.* were isolated from any of the 35 water and soil samples. The free chlorine level in potable water specimens was 1.0 parts per million.

Only 0.43 inches of rainfall were recorded for County A from May 1 through May 19, compared with a normal rainfall of 2.10 inches. At the peak of the outbreak, 1.2 inches of rainfall were recorded (Figure 1). The predominant wind direction during May (resultant wind vector) was from 241°,

TABLE 1—Attack Rates of Legionellosis in County A by Age*

Age Group (years)	No. of Cases	Attack Rate (per 10,000)
25–34	4	3.1
35–44	7	5.8
45–54	6	8.4
55–64	2	4.1
65–74	4	13.9
75–84	1	8.1

*Age unknown for 3 patients

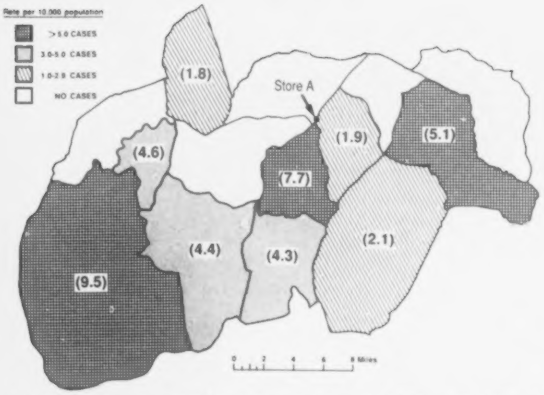

FIGURE 2—Census Tract of Residence for Persons with Legionnaires' Disease in County A, in Maryland, 1986 (numbers in parentheses indicate attack rate per 10,000 population; population per census tract ranged from 2,000 and 15,000)

TABLE 2—Antibody Titers to *Legionella pneumophila* among Employees of Selected Businesses

	Numbers of Employees with Indicated Titers					
Site of Employment	≤64	128	256	512	≥1024	Total
Store A	40	5	3	2	2	52
Store B	14	2	0	1	0	17
Near store A	33	3	1	0	0	37

from the southwest and toward store A from the construction site. Persons working in store A recalled extremely dusty conditions during May 1986, and review of the maintenance records for the ventilation system of store A revealed that frequent changes of the filters for the air intake had been required during the first three weeks of May. Store A was cooled with a freon-based, dry central air conditioning unit; there was no humidification of incoming air.

Discussion

Investigation of this Legionnaires' disease outbreak was complicated by logistic problems in obtaining and testing serum specimens that delayed laboratory confirmation of the diagnosis. We are unaware of previous Legionnaires' disease outbreaks occurring in rural areas despite serologic evidence that exposure is no less frequent in such areas.[7] Legionnaires' disease outbreaks in rural settings may be more difficult to identify because diagnostic facilities for legionellosis are frequently not available. Even in the context of epidemiologic study,[8] many patients with community-acquired pneumonia do not receive a precise etiologic diagnosis. Using review of radiograph reports to identify cases created additional difficulties in our investigation because of the low specificity of this method, unless combined with fever and white blood cell count.

Once Legionnaires' disease was identified as the cause of the outbreak, additional epidemiologic investigation demonstrated an association between visiting store A and illness. Results of the case-control study led to additional surveys to confirm the association of visiting store A and developing legionellosis. The attack rate of 1.5 percent in Store A employees is consistent with attack rates in other epidemics of Legionnaires' disease.[9,10] In addition, elevated antibody titers against *Legionella* were more common and these levels tended to be higher in employees of store A than in control populations.

Negative cultures of all environmental specimens, including water and soil samples, were not entirely unexpected since the specimens were obtained five weeks after the onset of the last case and the end of the outbreak. Since no evironmental legionellae were recovered, no microbiologic proof of the source of *Legionella* organisms was found in this outbreak. The absence of cooling towers at store A and the absence of exposure of case-patients to aerosols of potable water suggest that persons contracting Legionnaires' disease were exposed to some other source of *Legionnella* organisms.

Excavation and construction have been associated with outbreaks[10–12] and sporadic cases[13] of Legionnaires' disease, and excavation and construction of a shopping mall near store A occurred during the month of the epidemic. Although most outbreak investigations in which excavation and construction were associated with disease occurred before methods for recovering *Legionella* were available, in one nosocomial outbreak, *L. bozemanii* was recovered from potable water and from soil at a nearby excavation. *Legionella* was hypothesized to have colonized the potable water system at the time pipe connections linked the new building to the old.[11] Thus, in some outbreaks the link between excavation, construction, and Legionnaires' disease may be potable water.

In the current outbreak investigation cultures of potable water, although obtained five weeks after the end of the outbreak, were negative. High free-chlorine concentrations were found in the municipal water serving store A, so we doubt that *Legionella* could have survived in the potable water. Wind direction and the clogged air filters at store A suggest that airborne particles were blown from the construction site toward store A. Rains occurring on May 20, virtually the first rain of the month, coincided with a cessation of new cases after May 28, within the incubation period of Legionnaires' disease.

This temporal association is similar to that seen at the St. Elizabeth's Hospital outbreak in Washington, DC.[10] In that outbreak, attack rates of legionellosis for patients hospitalized in buildings near excavation were higher than the rates for patients not hospitalized near excavation. Because the etiologic agent was not discovered and methods to culture *Legionella* from the environment were not developed, the reservoir of *Legionella* was not determined. The tight temporal relationship between the onset of illness and excavation suggests that direct exposure to excavation was related to developing legionellosis.

Although legionellae have never been recovered from dry soil, they have been recovered from mud and wet soil.[11, 14–16] Since *Legionella* may be killed by heating or drying,[17] the hypothesis that legionellosis can occur after exposure to dry soil requires the assumption that *Legionella* may exist in an environment that provides protection from heating and drying. Additional environmental reservoirs for *Legionella* may be found in future studies, as were recently identified for *Blastomyces dermatitidis*.[18] The temporal relationship of the rains of May 20, and the cessation of new cases suggest that the rains halted the further transmission of *Legionella*.

This outbreak of Legionnaires' disease was associated with exposure to a retail store near a site of excavation and construction at a time when dry weather conditions prevailed. These results further suggest that, in some outbreaks, exposure to construction and excavation may be a risk factor for Legionnaires' disease. Future investigations of Legionnaires' disease outbreaks associated with construction and excavation must include expanded searches for environmental legionellae and consideration of novel reservoirs.

ACKNOWLEDGMENTS

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Dr. Robert S. Lawrence Named Editor of *American Journal of Preventive Medicine*

Robert S. Lawrence, MD, has been named editor of the *American Journal of Preventive Medicine*, a bimonthly publication sponsored by the American College of Preventive Medicine (ACPM) and the Association of Teachers of Preventive Medicine. Dr. Lawrence is Charles S. Davidson Associate Professor at Harvard Medical School and Chief of Medicine at Cambridge Hospital.

A fellow of ACPM and of the American College of Physicians, and a member of the Institute of Medicine, Dr. Lawrence recently served as chairman of the US Preventive Services Task Force, the national group whose work led to the landmark report on prevention in clinical practice—the *Guide to Clinical Preventive Services*.

A graduate of Harvard College and Harvard Medical School, Dr. Lawrence has also served in the Epidemic Intelligence Service (EIS) of the Centers for Disease Control. On the international level, he has worked as a medical epidemiologist in El Salvador, as coordinator of the Harvard-King Faisal University Project in Saudi Arabia, and as a consultant or visiting professor in Italy, Japan, and Qatar.

Dr. Lawrence serves on many national committees and has been especially concerned with human rights. He is author of various articles, book chapters and books with an emphasis on primary care, disease prevention, health promotion, and medical education.

The *American Journal of Preventive Medicine*, published by Oxford University Press, features original reviews, critical commentaries, editorials, book reviews, and correspondence on all aspects of practice, teaching, and research in preventive medicine. For futher information, contact ACPM, 1015 15th Street, NW, Suite 403, Washington, DC 20005; 202/789-0003.

Tuberculin Skin Testing in US Navy and Marine Corps Personnel and Recruits, 1980–86

ELEANOR R. CROSS, MS, AND KENNETH C. HYAMS, MD

Abstract: An extensive skin testing program is part of the United States Naval Medical Command's infectious disease control effort. From 1980 to 1986, 2,306,533 skin tests, using five TU PPD, were performed on active-duty Navy and Marine Corps personnel, 0.97 percent of which were positive. A downward trend in positive tests was found with a high of 1.43 percent in 1980 and a low of 0.80 percent in 1983. Since 1984, the percentage of positive tests has remained the same or increased. Shore-based medical facilities around the world reported 1,491,646 skin tests with 1.07 percent positive; Navy ships reported 814,887 skin tests with 0.78 percent positive. PPD-positivity for ships in the Pacific area was higher (0.98 percent) than for ships in the Atlantic (0.62 percent). During this same period, the percentage of positive tests in Navy and Marine Corps recruits ranged from a high of 1.82 percent in 1981 to a low of 1.23 percent in 1986. Since 1984, the percentage of positive tests has remained relatively stable in recruits. The frequency of positive PPD tests found in this study is lower than the percentage positive (1.59 percent) found in active-duty Navy personnel in 1969 and the percentage positive (5.2 percent) found in a study of Navy and Marine Corps recruits between 1958 and 1969. (*Am J Public Health* 1990; 80:435–438.)

Introduction

Since the early 1800s, the Medical Department of the US Navy has recognized the high transmission potential of tuberculosis.[1] Not only is there an increase in risk of exposure from world-wide travel but, as epidemiological studies have demonstrated, environmental conditions aboard ships are ideal for transmission.[2–4] Early attempts at TB control included screening sailors prior to an extended cruise and rejecting "all persons who were believed to be of unsound constitution."[1]

Today, the Navy's tuberculosis control program includes tuberculin skin test screening, contact investigation, and treatment of active cases and selected skin test reactors.[5] The Naval Medical Command directs all personnel whose previous test was negative to be periodically skin tested by the Mantoux method using 5 tuberculin units (TU) Purified Protein Derivative (PPD).[5] Annual screening is done on individuals assigned to operational units and to areas where possible exposure to tuberculosis is high, i.e., members of deployable Navy and Marine Corps units, shipboard personnel, and health care personnel. Sailors in low-risk areas, such as shore-based personnel in the United States, are tested every three years.

A person with induration measuring 10 mm or greater 48 to 72 hours after the administration of a 5 TU skin test is considered a "tuberculin reactor."[5] A new reactor is defined as "a person who has become a reactor within the past three years," and persons who are found to be reactors for the first time but whose prior tuberculin status is unknown. Since earlier studies indicated that a new reactor rate of 1 to 2 percent of personnel tested could be expected in many Navy and Marine Corps settings, the Naval Medical Command recommends a search for active cases of tuberculosis when the rate of new reactors exceeds 2.5 percent in groups of approximately 100 or more personnel.[5]

This report examines tuberculosis skin testing in the Navy from 1980 to 1986. Comparisons are made over time in and recruits Navy and Marine Corps active-duty personnel stationed in various geographic locations.

Methods

Each month, all outpatient facilities of the Naval Medical Command report the number of PPD tests performed and the number of reactors. Information on age, sex, race, and the exact size of the reaction is not recorded, however. In this study, computerized files of skin test results for the period January 1980 to December 1986 were analyzed.

Reports are received each month from approximately 600 shore-based facilities and from over 500 ships. Approximately 25 percent of the shore-based facilities are located outside of the United States. It is important to note that the location of the reporting facility may not necessarily be the location of the tuberculosis contact. Data from shore-based facilities may include skin test results on individuals from ships during port visits as well as results on individuals transferred to a facility from a different location.

Although these tests are administered and read by a number of different medical personnel, the Naval Medical Command provides very specific instructions on the administration and reading of these tests.[5] Therefore, the overall results of the skin tests from different locations are generally comparable. Tubersol (Connaught, Inc., Swiftwater, PA) was used for the skin testing during the period of this study.

For the 1980 to 1986 study period, the total number of skin tests performed in active-duty personnel fluctuated from a high of 373,747 in 1983 to a low of 287,251 in 1986. Average annual active-duty strength increased from 720,832 in 1980 to 805,128 in 1986. Consequently, approximately 40–50 percent of the total force was tested annually. The mean age of active-duty personnel increased from age 25.0 years in 1980 to age 26.1 years in 1986. The mean sex ratio of Navy and Marine Corps personnel changed from 16:1 (male:female) in 1980 to 11:1 in 1986.

For recruit personnel, approximately 83 percent were between the ages 17 to 21 years during the study period. The percentage of both Black and Hispanic recruits increased during the period 1981 to 1986 from 12 percent to 17 percent for Blacks and from 4 percent to 9 percent for Hispanics.

Results

Between 1980 and 1986, a total of 2,306,533 tests were read for all active-duty Navy and Marine Corps personnel, 0.97 percent of which were positive. The rate of positive tests decreased from a high of 1.43 percent in 1980 to a low of 0.80

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percent in 1983. In 1983, this downward trend changed, and from 1984 to 1986 the percentage of positive tests either remained the same or increased (Table 1, Figure 1).

Shore-based Facilities

In shore-based facilities located inside of the continental United States (CONUS), a total of 1,319,610 tuberculin skin tests were read, of which 1.00 percent were positive. The percentage of positive tests dropped from a high of 1.41 percent in 1980 to a low of 0.83 percent in 1983. After 1983, the rate of positive tests either rose or remained the same (Figure 2). The average percentage of positive tests in shore-based facilities located outside the continental United States (OCONUS) was higher (1.61 percent but followed a very similar trend (Figure 2).

By location of shore-based facilities, the highest frequency of positive tests was in the Philippines (3.10 percent), followed by Japan (2.90 percent), Korea (2.00 percent), Guam (1.88 percent), Bahrain (1.66 percent), and Okinawa (1.40 percent). By US Naval District, skin test positivity was highest in the Northwestern United States (1.79 percent) and Northern California (1.77 percent) (Table 1, Figure 3). The lowest rates were found in the North Central area (0.46 percent).

Positive TB Skin Tests Aboard Ships

Of 814,887 skin tests read aboard ships, 0.78 percent were positive. The positive rate was higher in ships based in the Pacific area (0.98 percent) than for ships based in the Atlantic (0.62 percent). As in shore-based facilities, there was a general downward trend in both areas from 1980 to 1982, especially in the Pacific. After 1982, the rates of positive tests remained fairly stable (Table 1, Figure 4).

Navy and Marine Corps Recruits

Between 1980 and 1986, a total of 618,074 skin tests were performed on recruits, 1.45 percent of which were positive (Table 1, Figure 5). The highest percentage of positive tests (1.82 percent) was found in 1981. Since 1984, the percentage of recruits with a positive PPD skin test has remained relatively stable at between 1.28 and 1.23 percent.

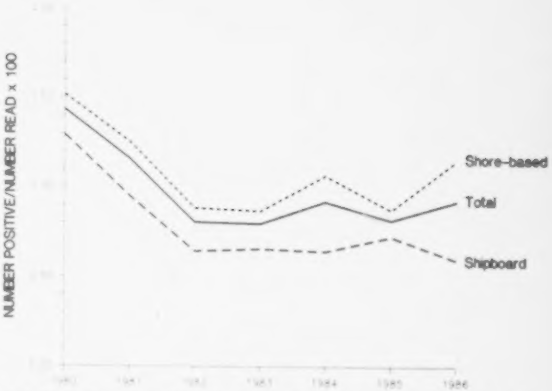

FIGURE 1—Frequency of Positive Tuberculin Tests in all Active-duty US Naval and Marine Corps Personnel, in all Shore-based Naval Medical Facilities, and in all ships, 1980–86

Discussion

The results of this study show a downward trend in the frequency of positive TB skin tests from 1980 to 1983. After 1983, the percentage of positive tests either remained stable or increased. It is noteworthy that these general trends occurred in all population groups of active-duty US Navy and Marine Corps personnel. The reason why different populations tested in different areas of the world exhibited such closely similar patterns of positive tests could not be determined. The data indicate, however, that the threat of TB infection is greatest in the Pacific for both shore-based and shipboard military personnel.

TABLE 1—Percent Positive Tuberculin Reactors in Active-duty US Navy and Marine Corps Personnel and in Recruits, 1980–86

	Number Reactors/Number Read × 100 $\frac{\text{Reactors}}{\text{Read}}$							
Location	Average 1980-86	1980	1981	1982	1983	1984	1985	1986
Total	0.97 $\frac{22388}{2308633}$	1.43 $\frac{4385}{305937}$	1.16 $\frac{3696}{318616}$	0.81 $\frac{2604}{323466}$	0.80 $\frac{2978}{373747}$	0.92 $\frac{3283}{356926}$	0.82 $\frac{2794}{340588}$	0.92 $\frac{2645}{287251}$
Shore-based								
Total	1.07 $\frac{16020}{1491646}$	1.52 $\frac{2908}{191658}$	1.25 $\frac{2750}{219710}$	0.88 $\frac{1927}{218103}$	0.87 $\frac{2158}{248741}$	1.07 $\frac{2401}{233717}$	0.87 $\frac{1847}{211232}$	1.15 $\frac{1939}{168465}$
CONUS	1.00 $\frac{13248}{1319610}$	1.41 $\frac{2363}{167518}$	1.18 $\frac{2355}{199209}$	0.84 $\frac{1622}{193872}$	0.83 $\frac{1860}{223659}$	1.00 $\frac{2088}{208074}$	0.78 $\frac{1402}{179945}$	1.06 $\frac{1578}{148433}$
OCONUS	1.61 $\frac{2772}{172036}$	2.26 $\frac{545}{24140}$	1.93 $\frac{395}{20501}$	1.26 $\frac{305}{24231}$	1.18 $\frac{298}{25182}$	1.53 $\frac{423}{27643}$	1.42 $\frac{445}{31287}$	1.89 $\frac{361}{19082}$
Atlantic	0.91 $\frac{420}{46396}$	1.33 $\frac{89}{6686}$	0.97 $\frac{51}{5245}$	0.89 $\frac{47}{5274}$	0.80 $\frac{44}{5526}$	1.26 $\frac{98}{7791}$	0.46 $\frac{51}{11089}$	0.83 $\frac{40}{4793}$
Pacific	1.87 $\frac{2352}{125640}$	2.61 $\frac{456}{17442}$	2.25 $\frac{344}{15256}$	1.36 $\frac{258}{18957}$	1.29 $\frac{254}{19656}$	1.64 $\frac{325}{19852}$	1.95 $\frac{394}{20218}$	2.25 $\frac{321}{14299}$
Shipboard								
Total	0.78 $\frac{6348}{814887}$	1.29 $\frac{1477}{114279}$	0.95 $\frac{939}{98908}$	0.64 $\frac{677}{105365}$	0.66 $\frac{820}{125006}$	0.64 $\frac{792}{123209}$	0.72 $\frac{937}{129356}$	0.59 $\frac{706}{118786}$
Atlantic	0.62 $\frac{2802}{454836}$	0.83 $\frac{478}{57459}$	0.77 $\frac{421}{54535}$	0.53 $\frac{305}{57699}$	0.52 $\frac{385}{73617}$	0.60 $\frac{420}{69837}$	0.66 $\frac{471}{71881}$	0.46 $\frac{322}{\text{[illegible]}}$
Pacific	0.98 $\frac{3546}{360051}$	1.76 $\frac{999}{56820}$	1.17 $\frac{518}{44371}$	0.78 $\frac{372}{47666}$	0.85 $\frac{435}{51389}$	0.70 $\frac{372}{53372}$	0.81 $\frac{466}{57475}$	0.78 $\frac{384}{\text{[illegible]}}$
Recruits	1.45 $\frac{8991}{618074}$	1.76 $\frac{1135}{64624}$	1.82 $\frac{1690}{92849}$	1.31 $\frac{1278}{97283}$	1.57 $\frac{1508}{95853}$	1.28 $\frac{1383}{108840}$	1.27 $\frac{1112}{87282}$	1.23 $\frac{907}{73723}$

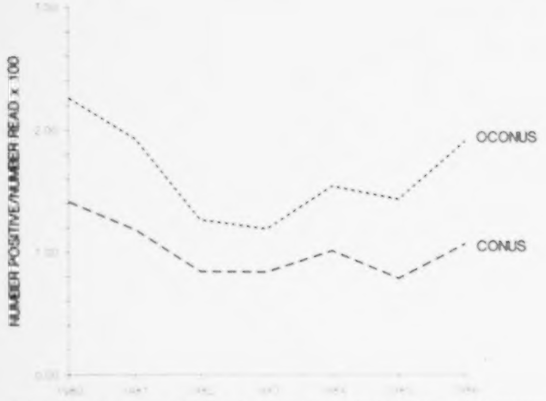

FIGURE 2—Frequency of Positive Tuberculin Tests in Active-duty US Naval and Marine Corps Personnel in Shore-based Facilities within the Continental United States (CONUS) and outside the Continental United States (OCONUS), 1980–86

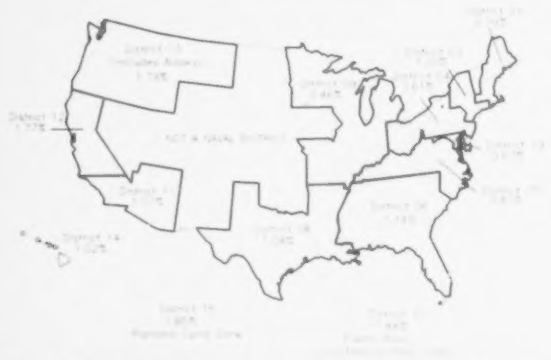

FIGURE 3—Percentage of Positive Tuberculin Tests in Active-duty US Naval and Marine Corps Personnel by US Naval District, 1980–86

In 1969, 1.59 percent of skin tests were positive overall; 1.72 percent were positive in shore-based facilities in the continental United States; and, 1.08 percent of PPD tests were positive on Navy ships.[6] This level of skin test positivity is higher than for comparable groups during the 1980 to 1986 period of this study and suggests a continuation of the downward trend of tuberculosis infection in the US Navy which was observed prior to 1970.[7]

Between 1958 and 1969, Navy recruits were skin tested with 5 TU PPD by specially trained public health nurses.[7] Tuberculin reactions measuring 10 mm or greater were found in 5.2 percent of the tests. This is substantially higher than the percentage of positive tests (1.45 percent) found in recruits from 1980 to 1986 and corresponds to the lower incidence of tuberculosis in the United States during the period of the current study.

One problem when comparing positive skin test rates in military personnel is that an unknown number of tests are applied but not read, which would lower overall rates. However, there is no reason for this problem to have been greater in any location or any particular year of this study. Also, a person with a positive skin test who is missed on one screening has another chance of being detected on subsequent testing and counted as PPD-positive.

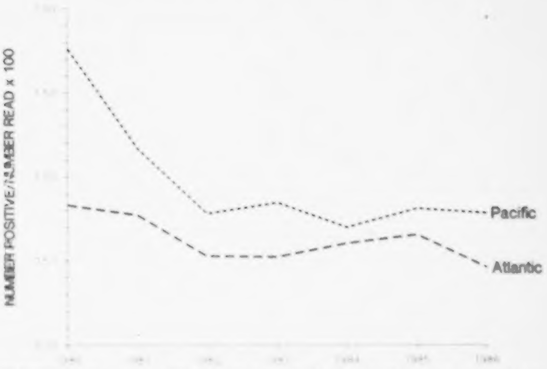

FIGURE 4—Frequency of Positive Tuberculin Tests in Active-duty US Naval and Marine Corps Personnel Assigned to Ships in the Atlantic and the Pacific, 1980–86

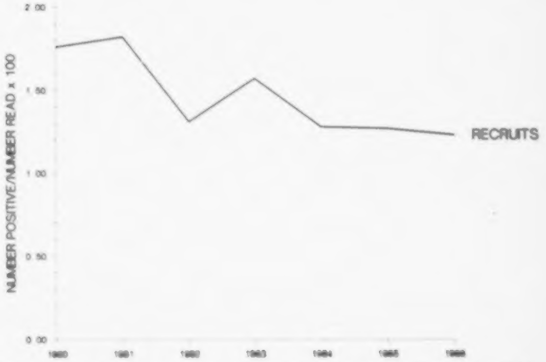

FIGURE 5—Frequency of Positive Tuberculin Tests in US Naval and Marine Corps Recruits, 1980–86

There was a steadily declining incidence rate of tuberculosis disease in the US civilian population until 1984. In 1986, for the first time since 1953, there was a substantial increase in indigenous tuberculosis cases.[8] This reversal has been attributed to an increase in the number of cases of acquired immunodeficiency syndrome (AIDS); to an increase in the number of tuberculosis cases in the homeless, immigrant and other susceptible populations; and, to decreased tuberculosis control efforts.[8-10]

The military population in this study also demonstrated a change in the declining incidence of tuberculosis infection at the time tuberculosis disease was increasing in the civilian population. The military population, however, does not include a large number of individuals who are particularly susceptible to tuberculosis infection, such as the homeless or immigrants, although AIDS cases have been reported in the military.[11,12] Also, the highest rates of positive tests observed in this study occurred in shore-based facilities outside the continental United States, where events occurring in the US may have had less impact.

Factors which could have changed the extent of TB exposure in the military, such as increased deployment to and increased recruitment from endemic regions such as the Philippines, could not be analyzed from the available data.[7,13,14] However, no large shipboard tuberculosis outbreaks of the magnitude previously reported occurred during this seven-year study period.[2,3]

The "booster effect" is a phenomenon in which increases in the size of the skin test reaction is due to boosting of waned hypersensitivity by sequential testing rather than new infection. Boosting is more common among middle-aged and older persons and when skin testing is done at frequent intervals.[15,16] Because the Naval Medical Command recommends annual testing for personnel in high risk areas and retesting every three years for those personnel at low risk, boosting should be uncommon except in frequently tested groups such as hospital personnel.

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5th National Conference on Chronic Disease Prevention and Control

The Centers for Disease Control, the Association of State and Territorial Health Officials (ASTHO), and the Directors of State Chronic Disease Programs will cosponsor the fifth national conference on "Chronic Disease Prevention and Control: From 1990 to 2000," October 17–19, 1990 at the Westin Hotel, Renaissance Center, in Detroit, Michigan. The conference is open to the public; there will be a $25 registration fee.

The 1990 conference will include two plenary sessions that address demographic trends:

- Implications for Chronic Disease and the Public's Health
- Tobacco Control: Science and Politics

Six concurrent sessions of invited papers will focus on special populations:

- Minorities, the Elderly, and Rural
- State Highlights—A Look at Michigan
- Quality of Life and Disability—Issues and Impact
- Dietary Fat/Nutrition
- Industry and Chronic Disease Prevention—What Are the Societal Roles and Responsibilities?
- State and Community Tobacco Control Activities

The conference will consider abstracts for oral presentations, poster sessions, or roundtable discussions on relevant chronic disease topics involving: program development and implementation; epidemiologic studies; and social and economic issues.

For additional information, contact: Jack Friel, MPA, Chief, Conference Management Operations, Center for Chronic Disease Prevention and Health Promotion, Mailstop A37, Centers for Disease Control, Atlanta, GA 30333; (404) 639-2249 FAX (404) 639-1563.

Tuberculosis Infection in Urban Adolescents: Results of a School-based Testing Program

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Sue Weidhaas Etkind, RN, MS, Carol Almeida, BSN, John Bernardo, MD, and George A. Lamb, MD

Abstract: In 1986–87 a pilot tuberculosis (TB) skin testing program was introduced for seventh and tenth grade students in the Boston (Massachusetts) public schools. The 8.9 percent tuberculin positivity rate in tenth grade students was significantly higher than the 5.1 percent rate found in seventh graders. A majority of those who were skin test positive were born outside the United States. These results suggest that tuberculin testing in an urban school setting may identify a significant number of candidates for TB preventive therapy, particularly among tenth grade students and those who are foreign-born. (*Am J Public Health* 1990; 80:439–441.)

Introduction

Tuberculosis (TB) continues to be an important public health problem in the United States. In 1986, 22,768 cases of tuberculosis were reported nationwide, including 1,214 (5.3 percent) in children younger than 15 years of age.[1] In 1987, nine of 123 verified TB cases in Boston, Massachusetts occurred in children under age 15.[2]

Development of active TB can be largely prevented by isoniazid (INH) therapy in infected persons without disease.[3] However, identification of both adults and children with TB infection is problematic. Current guidelines from the American Academy of Pediatrics for TB testing in children advise annual tuberculin testing in high-risk children, such as those in neighborhoods where the TB case rate is higher than the national average, those from households with one or more cases of TB, American Indian and Alaskan children, and children of parents who have immigrated from Asia, Africa, the Middle East, Latin America, or the Caribbean.[4] In low-risk groups, TB skin testing is recommended at 12–15 months of age, before school entry, and in adolescence. However, no recent published data on the prevalence of tuberculous infection in urban adolescents in the United States are available to assess whether adolescents should be considered at high risk.

Following the diagnosis of TB in a Boston Public School student in the fall of 1986, a TB testing program was introduced for seventh and tenth grade students. Students with positive skin tests were referred for appropriate medical evaluation and preventive therapy. The purpose of this report is to describe the prevalence of tuberculin skin test positivity in seventh and tenth grade students who participated in the program as well as the subsequent course of preventive therapy in those for whom such therapy was indicated.

Methods

The Boston Public School system enrolls approximately 57,000 students. Seventh graders are enrolled in 25 middle schools, and tenth graders in 19 high schools with each school assigned either a full- or part-time nurse. Racial-ethnic breakdown reveals that 47.8 percent of all Boston Public School students are Black, 25.6 percent White, 17.7 percent Hispanic, 8.2 percent Asian, and <1 percent other.

At the time this program was introduced, TB testing was required for new school entrants only. The new policy continued to require TB testing for new school entrants (regardless of grade). In addition, seventh and tenth grade students not previously known to be PPD (purified protein derivative) positive were required to have a TB skin test done during that school year. Due to problems of retention and motivation of urban adolescents in school, however, students were not excluded from classes if they did not comply with the seventh and tenth grade testing policy. Students were encouraged to have tuberculin testing performed by their primary care provider, but arrangements for placement and reading of skin tests in the schools were made. Mantoux testing using intermediate strength PPD was the preferred method for skin testing as suggested by the Centers for Disease Control (CDC) and the American Thoracic Society guidelines.[5] Tubersol (Connaught Laboratories Inc., Swiftwater, PA) was the only PPD product used within the schools. Medical evaluation and a single chest x-ray following a documented positive skin test were mandatory.

Data collected for each tuberculin positive student included information on age, sex, race, country of origin, site of testing, size of PPD induration, site of medical follow-up, and duration of INH therapy. For analysis, chi square testing was used (Professional Database Analysis System; Conceptual Software, Inc., Houston, Texas).

Results

Of 3,642 seventh graders eligible, 2,799 (77 percent) were tested (Table 1). In the tenth grade 2,584/4,353 (59 percent)

TABLE 1—Results of a Tuberculosis (TB) Testing Program, Boston Public Schools, 1986–87

	Grade 7	Grade 10
Number eligible	3,642*	4,353*
Number (percent) tested	2,799 (77)	2,584 (59)
Number (percent)** PPD positive***	145 (5.1)#	233 (8.9)

*This number excludes 34 and 29 known PPD positives in seventh and tenth grades, respectively.

**Numbers include those previously known to be PPD positive.

***PPD positive is defined as induration ≥ 10 mm at 48–72 hours.

#Difference in proportions is 3.8 percent (95% CI = 2.5, 5.2) using test-based confidence interval.

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participated in the program. Participation rates in individual schools ranged from 33 percent to 100 percent in seventh grades and from 16 percent to 100 percent in tenth grades. Of seventh grade program participants, 1,554 (56 percent) were tested in school and 1,245 (44 percent) by their private health care provider. Of tenth graders, 1,881 (73 percent) were tested in school and 703 (27 percent) by a private provider. The rate of tuberculin positivity in seventh graders (5.1 percent) was substantially lower than that in tenth graders (8.9 percent) (see Table 1).

Mean age of seventh graders who tested positive during the program was 13.1 years (range 12 to 15 years); 85 percent were foreign-born. The mean age of tuberculin positive tenth graders was 16.5 years (range 14 to 21 years); 80 percent were foreign-born (Table 2). Country of origin was documented for 295 tuberculin positive students (Table 3). The largest number were from Caribbean countries with Haiti the most common country of origin.

Information on the size of PPD induration was available for 131 of the 315 students with induration ≥ 10 mm. Twenty-four percent had a reading ≥ 20 mm; an additional 13 percent had induration between 16 and 19 mm.

Records to assess compliance with preventive therapy were available for 230 of 314 (73 percent) tuberculin positive students (Table 4). Overall 63 percent of students for whom preventive therapy was recommended completed at least six months of therapy.

An unexpected finding during the program was the discovery of an unsuspected case of active pulmonary TB. The individual was an asymptomatic 20-year-old Haitian male in the tenth grade whose tuberculin test had been negative two years previously. His chest x-ray, done in follow-up to a 25 mm PPD in school, revealed right upper lobe infiltration consistent with active or old TB. Sputum smear was negative. Isoniazid, rifampin, and ethambutol therapy was started. When cultures revealed *M. tuberculosis* sensitive to all antimicrobial agents tested, ethambutol was discontinued. In the high school he attended 41 percent of the tenth grade students participated in the program. Thirty-three (18 percent) of these were tuberculin positive.

TABLE 2—Demographic Characteristics of Students with a Positive* Tuberculin Skin Test during a School-based Testing Program, 1986–87

Characteristics	Grade 7 (N = 111)	Grade 10 (N = 203)
Mean age (years)	13.1	16.5
Sex**		
Male	2.7	5.4
Female	2.5	4.9
Race**		
White	.46	1.4
Black	1.4	4.9
Hispanic	3.2	6.4
Asian	13.8	13.5
Place of Birth		
Foreign-born	95 (85)***	163 (80)
US-born	8 (7)	29 (14)
Unknown	8 (7)	11 (5)

*Positive is defined as induration ≥ 10 mm at 48–72 hours using Mantoux testing.

**Numbers expressed as rate/100 using the total number of students in that category in each grade as denominator.

***Numbers in parentheses denote percentage. May not add up to 100 due to rounding error.

TABLE 3—Birthplace of Tuberculin Positive Students, 1986–87 (N = 295)

Region	Number (%)
Caribbean	
Haiti	51 (17)
Dominican Republic	20 (7)
Puerto Rico	11 (4)
Other	31 (11)
Asia	
China	43 (14)
Vietnam	29 (10)
Cambodia	10 (3)
Other	5 (2)
North America	
United States	37 (13)
Central America	
Other	21 (7)
Africa	
Cape Verde	19 (6)
Other	5 (2)
South America	8 (3)
Europe	5 (2)

*Percentages do not add up to 100 due to rounding error.

TABLE 4—Therapy of Tuberculin Positive Boston Public School students, 1985–87

Results	Grade 7 (N = 111)	Grade 10 (N = 203)	Total (N = 314)
No record	37	40	77
Moved	1	6	7
Records reviewed	73	157	230
Previous Rx	7 (10)*	12 (8)	19 (8)
Rx < 6 mos**	19 (26)	65 (41)	84 (37)
Rx ≥ 6 mos	47	80	127 (55)

*Percentage of those in whom record was available for review.

**Includes five students in whom therapy had to be stopped due to adverse effects.

Discussion

In 1979, data from 13 cities with populations of 250,000 or greater revealed a tuberculin positivity prevalence of 5.6 percent in those ages 5–14 years.[6] In our study, the 8.9 percent rate of tuberculin positivity found in tenth graders whose mean age was 16.5 years was substantially higher than the 5.1 percent found in seventh graders whose mean age was 13.1 years. There are several possible explanations for the higher rate in tenth graders. There may be proportionately more foreign born tenth graders than seventh graders; data to document this possibility are not available. Second, the higher rate may be related to more advanced age and consequently a longer chance to become infected. Third, higher rates in tenth graders may have occurred as a result of spread from the tenth grade student with unsuspected Class III pulmonary TB. However, the case was not smear positive, and excluding results from his school lowers the prevalence rate in tenth graders only to 8.3 percent (200/2405).

Generalizability of tuberculin prevalence rates noted in this study to all urban adolescents is limited by several factors. Rates in students may differ from those in adolescent drop-outs and student non-participants. Although data are available for 59 percent of tenth graders and 77 percent of seventh graders, little is known about those who chose not to participate. Finally, Boston is a gateway city for refugees; this may result in higher rates of tuberculin positivity than in cities where refugees are less common.

Delivery of adequate preventive therapy, defined as ≥ six months of INH, could be documented in a greater proportion of seventh graders than tenth graders (74 vs 59 percent). These figures are comparable to the 73.4 percent completion of preventive therapy figure noted by CDC among all persons for whom preventive therapy was recommended over a 12-year period in 16 selected regions of the US.[7]

Current TB surveillance practices in the United States vary widely. A 1983 survey of health department officials revealed that only 20 percent of respondents required TB screening in students.[8] The authors concluded that testing of students was non-productive other than "in schools with large numbers of refugees, immigrants, children of indigent parents, etc." The proportion of urban school systems which meet such criteria is unknown, and data from other geographic regions are needed to resolve the issue.

In the United States, TB rates are higher in minorities. In 1986, the TB case rate per 100,000 population was 4.6 in Whites, 17.7 in Hispanics, 27.6 in Blacks, and 50.4 in Asian/Pacific Islanders.[9] Our tuberculin positivity rates reflect this trend with the highest prevalence in Asians, followed by Hispanics and Blacks.

Using data from an analysis done by the Centers for Disease Control in 1986,[10] the cost effectiveness of this program can be demonstrated. Assuming progression to active TB in 10 percent of those not receiving preventive therapy, and $4,983 cost for a single uncomplicated case of disease, the cost of preventable cases detected in this program would be $189,384. In contrast, the actual cost of the screening was $75,465 ($19,000 for a part-time nurse coordinator, $1,057 for screening supplies, and $55,408 for the clinical and pharmaceutical costs related to INH follow-up) for a net savings of $113,919.

Preventable TB cases continue to occur in children in the US.[11] Our data suggest that TB testing in urban school settings comparable to ours is likely to uncover a significant number of adolescents who are tuberculin positive. Because of the relatively high risk of progression to active tuberculosis in adolescents who are infected with TB, identification of and therapy for tuberculin positive children and adolescents is particularly important.[12]

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AAFP Will Publish Family Practice Research Journal

Family Practice Research Journal is now an official journal of the American Academy of Family Physicians (AAFP), signifying the Academy's efforts to support clinical research by family physicians. The Academy recently assumed ownership from Sciences Press, a subsidiary of Plenum Publishing Corporation, New York. *American Family Physician* also is a medical journal of the Academy.

According to the publisher, Clayton Hasser, "*Family Practice Research Journal* aims to complement other family journals by focusing on the stimulation and renovation of clinically oriented, practice-initiated and practice-based research articles, with emphasis on assisting novice family physician researchers and authors to implement and publish their research studies." Currently published twice a year, the journal will become a quarterly publication beginning with the June 1990 issue.

The American Academy of Family Physicians, headquartered at 8880 Ward Parkway, Kansas City, MO 64114, is the national association of family doctors. It is the largest national medical specialty organization with more than 68,000 members. The Academy was founded in 1947 to provide high continuing medical education standards.

Epidemiologic and Clinical Characteristics of Persons with Serologic Evidence of *E. canis* Infection

Barton W. Rohrbach, VMD, MPH, John R. Harkess, MD, S. A. Ewing, DVM, PhD,
Jerry Kudlac, MPH, Garry L. McKee, PhD, and Gregory R. Istre, MD

Abstract: The results of a serosurvey of Oklahomans for the presence of antibody to *Ehrlichia canis* is reported. Paired serum specimens, from patients lacking the serologic criteria for diagnosis of Rocky Mountain spotted fever (RMSF), were tested. A four-fold increase in *E. canis*-IFA antibody was found in 16/144 (11 percent) of these paired serum samples. Patients with serologic evidence of *E. canis* infection had a mean age of 34 years, 69 percent were male, and 63 percent lived in a town <10,000 population. Signs and symptoms included: fever 94 percent, headache 94 percent, fatigue 94 percent, anorexia 81 percent, nausea 60 percent, and rash 44 percent. When compared to control patients, whose sera were submitted for RMSF testing but did not meet serologic criteria for RMSF or *E. canis*, case-patients were more likely to have had leukopenia (OR = 4.9, 95 percent CI = 1.2, 19.0) and tick exposure (OR = 9.5, 95 percent CI = 1.4, 62.7). The results suggest *E. canis*, or a closely related agent, is a cause of human illness. Ticks are a probable vector. (*Am J Public Health* 1990; 80:442–445.)

Introduction

In each of the years 1982–86, Oklahoma has had the highest rate of Rocky Mountain spotted fever (RMSF) in the United States. For every case of RMSF reported, at least four other persons were tested for RMSF, but were negative for RMSF antibody.[1] We have suspected that these illnesses, which we have called "Oklahoma tick fever," may be caused by other tick-borne agents.[1]

Ehrlichia canis is the causative agent of canine ehrlichiosis, a tick-borne, rickettsial disease of dogs with wide geographic distribution.[2,3] In 1986, the first case of human infection with *E. canis* was described.[4] The patient was bitten by several ticks 12 days before onset of illness and was initially suspected to have RMSF. The diagnosis of *E. canis* infection was made from electron micrographs of leukocytic cytoplasmic inclusions characteristic of *Ehrlichia* species, and the presence of an eight-fold fall in serum antibody titer to *E. canis* during the seven weeks following onset of illness. Since discovery of the first case, additional patients have been identified, by serologic testing, from at least 11 states.[5–9] Although *Ehrlichia canis* has not been isolated under natural conditions from species other than *Canidae*, current evidence supports the role of *E. canis*, or a closely related agent, as a cause of illness among human beings.[2]

We undertook a case-control study to link clinical characteristics and risk factors with serologic evidence of *E. canis* infection.

Methods

Case Surveillance

To identify patients with serologic evidence of *E. canis* infection, we tested paired (acute- and convalescent-phase) sera from 144 persons. The survey included all available paired serum specimens that were tested for *R. rickettsii* in Oklahoma during 1986 and found negative. We obtained sera from the Oklahoma State Department of Health (OSDH) and a commercial hospital laboratory located in Oklahoma City. Using an indirect fluorescent antibody test (IFA), the convalescent-phase specimen for each pair was screened for *E. canis* antibody at 1:16 dilution. For patients whose sera was positive at 1:16, the acute- and convalescent-phase specimens were tested in consecutive two-fold dilutions until an end point titer was reached. A positive test result was defined as a four-fold or greater increase in antibody titer to *E. canis* antigen, and a peak convalescent-phase titer of ≥1:64.

Sera from patients with positive tests for *E. canis* were also tested for antibodies to *Rickettsia typhi* and *Coxiella burnetti* by microagglutination test at the Rocky Mountain Laboratories, Hamilton, Montana.[10] All available blood smears from case-patients, made during their illness, were examined for leukocytic intracytoplasmic inclusions. Blood specimens obtained from dogs belonging to case-patient households were tested for *E. canis* antibody at the Oklahoma Animal Disease Diagnostic Laboratory, Oklahoma State University, Stillwater.[11] Three ten-fold serum dilutions beginning at 1:10 were tested for each dog specimen.

Case-Control Study

We compared the epidemiologic and clinical characteristics of the 16 patients with positive tests for *E. canis* with 32 controls randomly selected from the 128 RMSF-IFA negative patient serum pairs remaining after case-patients had been identified. Controls were matched by laboratory from which the case-specimen originated, and consisted of patients whose acute- and convalescent-phase serum was <1:64 for *R. rickettsia* and convalescent-phase serum <1:16 for *E. canis*.

Efforts were made to conduct personal interviews for case-patients and controls. Substitute control-patients were randomly chosen from the remaining controls when it was not possible to complete interviews with the initially selected controls. Interviews began with open-end questions and concluded with questions designed to elicit specific clinical and demographic information and animal and arthropod exposure histories. All available medical records were reviewed for cases and controls. Odds ratios (OR) and 95%

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confidence intervals (CI) were calculated using the program developed by Rothman and Boice.[12]

Laboratory Procedure

A standard IFA procedure for detecting antibody to *R. rickettsii* had been used by both laboratories from which RMSF-negative serum was obtained.[13] All testing of human serum for *E. canis* antibody was done at the Oklahoma State Department of Health (OSDH) laboratory using a commercially available *E. canis* antigen (ProtaTek International Inc., St. Paul, Minnesota*), and according to a published procedure[11] with the following exceptions: human serum was substituted for canine serum, dilutions started at 1:16, a quantitative control was used, and goat anti-human conjugate (Clinical Sciences, Wippany, New Jersey*) was substituted for anti-dog conjugate. A random sample of case-patient sera was blindly retested for *E. canis* antibody by OSDH laboratory personnel. Agreement was confirmed if the second test result was within one dilution of the first result.

Results

Case Surveillance

Of 144 serum pairs tested, 16 (11 percent) had a four-fold increase in antibody titer to *E. canis* antigen (Table 1). The geometric mean titer for case-patients increased during each successive week for the first month after onset. An agreement of 80 percent between the first and second test result was recorded when 10 of the serum specimens were blindly retested for *E. canis* antibody.

Case Description

The mean age of case-patients was 34 years (range 2–68), and more than two-thirds were male (Table 2). Most case-patients lived in a town of less than 10,000 population or a rural unincorporated area, and all but one case-patient recalled finding one or more ticks on their body during the 30 days before onset of illness.

TABLE 2—Demographic Characteristics and Exposures for Patients with Serologic Evidence of *E. canis* infection, Oklahoma, 1986

	Cases No./Total (%)	Controls† No./Total (%)	Odds Ratio	95% Confidence Interval
Age X = 34 (range 2-68) X = 28 (range 2-80)				
Sex (male)	11/16(69)	19/32(59)	1.5	(0.4, 5.4)
Reside in town <10,000	10/16(63)	13/32(41)	2.4	(0.7, 8.4)
Arthropod exposure,				
Ticks on body				
Attached	12/16(75)	16/31(52)	2.8	(0.6, 13.3)
Attached or crawling*	15/16(94)	19/31(61)	9.5	(1.4, 62.7)
Chiggers	7/16(44)	10/30(33)	1.6	(0.4, 5.5)
Fleas	0/16(00)	9/30(30)		
Mosquitoes	8/16(50)	18/30(60)	0.7	(2.3, 0.2)
Dog contact	14/16(88)	23/31(74)	2.4	(0.5, 12.9)
Two or more dogs	6/16(38)	6/31(19)	2.5	(0.7, 9.6)
Other animals	5/16(31)	10/29(34)	0.9	(3.2, 0.2)

†Denominator may be less than 32 due to missing information.

Signs and symptoms most frequently experienced by case-patients included fever, headache, fatigue, anorexia, and nausea (Table 3). Rash was present in 7/16 (44 percent) patients, and was described as petechial (three patients), maculopapular (three patients), or vesicular (one patient). When present, rash was limited to the trunk and proximal extremities; no case-patient had rash involving hands or feet.

Results of 43 total leukocyte counts performed on 15/16 (94 percent) case-patients were evaluated. Leukopenia, (WBC $<4{,}800/\text{mm}^3$), was observed in 10/15 (67 percent) case-patients. Excluding patient #4, whose onset date is uncertain, the median day after onset of clinical signs for the lowest total leukocyte count was day 4 (range 0–11 days) for the remaining nine case-patients who had leukopenia. Of 35 total lymphocyte counts available for 14 case-patients, lymphopenia, ($<1{,}500/\text{mm}^3$), occurred in 9/14 (64 percent), and

TABLE 1—Laboratory Findings for Patients with Serologic Evidence of *E. canis* Infection, Oklahoma, 1986

Patient Number	Age (years)	Sex	Total			AST† (SI)	Titer 1	Titer 2	Titer 3
			WBC*	Lymphocytes	Platelets*				
1	02	M	1,600	800	31,000	23.50	<16 (03)**	2,560 (19)**	128 (216)**
2	59	M	2,500	1025	NORMAL	–	16 (03)	64 (22)	128 (189)
3	59	F	2,100	630	68,000	0.48	16 (14)	64 (22)	32 (93)
4	46	M	4,000	1040	357,000	0.70	<16 (84)	64 (99)	128 (140)
5	36	F	6,300	1827	–	–	<16 (01)	64 (13)	128 (204)
6	10	F	4,500	2610	NORMAL	–	32 (04)	128 (12)	128 (69)
7	28	M	9,700	1455	275,000	–	<16 (03)	64 (34)	128 (121)
8††	12	M	3,400	1836	173,000	11.50	<16 (02)	128 (16)	128 (116)
9	40	M	1,600	464	71,000	3.43	16 (04)	256 (10)	64 (278)
10	09	M	–	–	–	–	16 (03)	128 (18)	32 (233)
11††	39	M	10,400	2496	NORMAL	–	16 (15)	256 (29)	32 (152)
12	07	M	2,800	1456	–	–	32 (03)	128 (18)	32 (225)
13	66	F	1,900	570	33,000	6.45	16 (05)	128 (24)	256 (207)
14	37	M	10,900	6758	153,000	–	32 (15)	128 (29)	– –
15	26	M	5,500	2214	–	–	64 (15)	256 (24)	128 (236)
16	68	F	4,300	645	380,000	0.48	32 (01)	64 (11)	128 (128)

*Lowest value recorded.
†Highest value recorded, reference range 0–0.58 ukat/L.
**Days after onset of illness.
††Patient had other laboratory confirmed diagnosis.

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*Use of trade names is for identification only and does not imply endorsement by the United States Department of Health and Human Services, or the United States Public Health Service.

TABLE 3—Clinical Characteristics of Patients with Serologic Evidence of *E. canis* Infection, Oklahoma, 1986

	Cases* No./Total (%)	Controls† No./Total (%)	Odds Ratio	95% Confidence Interval
Signs and Symptoms				
Fever	15/16(94)	28/32 (88)	2.1	(0.2, 20.5)
Headache	15/16 (94)	22/32 (69)	6.8	(1.0, 48.3)
Fatigue	15/16 (94)	30/32 (94)	1.0	(1.0, 1.0)
Anorexia	13/16 (81)	24/30 (80)	1.1	(0.2, 5.1)
Nausea	9/15 (60)	19/32 (59)	1.0	(0.3, 3.5)
Myalgia	7/15 (47)	17/32 (53)	0.8	(0.2 2.7)
Vomiting	7/16 (44)	17/32 (53)	0.7	(0.2, 2.3)
Rash	7/16 (44)	22/32 (70)	0.4	(0.1, 1.2)
Adenopathy	3/16 (19)	6/32 (19)	1.0	(1.0, 1.0)
Abdominal pain	3/16 (19)	6/32 (19)	1.0	(1.0, 1.0)
Confusion	2/16 (12)	4/32 (12)	1.0	(1.0, 1.0)
Hospitalized	6/16 (38)	10/32 (31)	1.3	(0.4, 4.7)
Treatment with chloramphenicol or tetracycline	14/16 (94)	28/31 (88)	0.8	(0.1, 12.3)
WBC <4,800	10/15 (67)	7/24 (29)	4.9	(1.2, 19.0)
Lymphocytes <1,500	9/14 (64)	11/23 (48)	2.0	(0.4, 9.7)
Platelets <150,000	5/13 (38)	4/18 (22)	2.2	(0.4, 14.1)

*Denominator may be less than 16 due to missing information.
†Denominator may be less than 32 due to missing information.

excluding patient #4, the median day after onset of illness for the lowest lymphocyte count was 6.5 (range 0–15 days). Platelets were evaluated in 13 case-patients, five (38 percent) of which were thrombocytopenic, (<150,000/mm^3). A total of 21 platelet counts were performed on the five thrombocytopenic case-patients, and the median day after onset for the lowest platelet count was 6 (range 3–16 days).

Of seven case-patients for whom data were available, five (71 percent), had serum aspartate aminotransferase (AST) values above the reference range (0–0.58 ukat/L). Serum from six of these patients was also tested for total bilirubin, and in two patients (Nos. 8,9 Table 1) the values were 6.6 and 2.2 mg/dL, respectively.

Sufficient paired serum was available from 13/16 (81 percent) case-patients to test for antibodies to *Rickettsia typhi* and *Coxiella burnetti*. No significant titers (>1:8) were detected to either antigen. Microscopic examination of peripheral blood smears from three case-patients and a buffy coat smear from a fourth, made during clinical illness, did not reveal inclusions characteristic of *Ehrlichia* spp.

Illness resulted in hospitalization for 6/16 (37 percent) case-patients, and tetracycline or chloramphenicol was prescribed for all but two. There were no case-patient deaths.

Two case-patients were diagnosed with other illness. One 12-year-old boy (No. 8) had vomiting, jaundice, petechial rash, and a positive IgM-specific anti-HAV test diagnostic of recent hepatitis A infection; the other, a 39-year-old male (No. 11), had fever, fatigue, posterior cervical adenopathy, and a four-fold decrease in antibody titer to Epstein-Barr virus supporting a diagnosis of mononucleosis. A third case-patient (No. 4) developed a chronic illness with fatigue, myalgia, and a vesicular rash on his arms, head and mouth two to four weeks after removing a tick from his abdomen. One month after onset he received a 10-day course of tetracycline 250 mg QID. Serum specimens drawn 84, 99, and 140 days after onset revealed no antibodies to *R. rickettsii* but elevated titers to *E. canis*. The cause of illness in the remaining 13 case-patients was either unknown or clinically thought to be RMSF despite negative serologic tests.

Of 25 dogs in 14 case-patient households, we tested 16 (64 percent), from nine households, for *E. canis* infection and none had antibody titers, ≥1:100, suggestive of recent infection.

Case Control Study

All cases, and 27/32 (84 percent) controls were interviewed in person by one or two of the authors (BWR and JRH). Cases and controls were informed of serologic results prior to interview. Of 32 original controls, two were deceased, two had moved out of state, and one could not be contacted. Information for the two deceased controls was obtained from medical records, and telephone interviews were conducted for three others, two of whom had moved out of state and one for whom personal interview was impractical due to constraints of time and distance.

The median interval of time from onset of disease to collection of acute- and convalescent-phase serum was similar for cases (median = 4 and 18.5 days, respectively), and controls (median = 3 and 17 days, respectively). Case-patients were slightly older, more often male, and somewhat more likely to live in a town of <10,000 population or a rural unincorporated area than were controls (Table 2).

Signs and symptoms in case-patients were similar to those of controls except for an increased frequency of headache reported by case-patients. Case-patients were more likely to have developed leukopenia (WBC <4,800/mm^3) and to have reported finding a tick on their body in the month before onset of illness both of which were statistically significant (Table 2). Exposure to bloodsucking arthropods other than ticks was similar for case-patients and controls except that controls were more likely to report flea bites.

Discussion

Our findings support the role of *E. canis*, or an antigenically similar agent, as a cause of clinical illness among humans in Oklahoma. We have demonstrated serologic evidence of *E. canis* infection in a subgroup of Oklahomans whose serum was submitted for RMSF testing, but that did not meet the criteria for serologic confirmation. Laboratory findings of leukopenia, lymphopenia, and elevated hepatic enzymes in both canine [14] and human[4–9] patients are consistent with the hypothesis of a common agent. The similarity of clinical characteristics among case-patients and the differences between case-patients and controls support the hypothesis that the case-patient syndrome is a distinct clinical entity.

E. canis is reported to have a limited host range,[2] and the only recognized vector, *Rhipicephalus sanguineus*, feeds infrequently on humans.[15] This suggests the possibility of an antigenically similar, but distinct, agent as the cause of illness in case-patients, or an unrecognized mechanism of transmission for *E. canis*.

The proportion of *E. canis* positive serum pairs in the sample tested (16/144, 11 percent) when applied to the total number of serum specimens found negative for RMSF in Oklahoma each year implies that the incidence of human ehrlichiosis may be as much as half that of RMSF. The frequency with which evidence of infection was found in Oklahomans and the wide distribution of *E. canis* in the United States suggests that other states should consider surveillance to determine the presence and distribution of infection.

Further studies are necessary to assess the sensitivity, specificity, and reliability of the canine IFA test for humans. Our results, on a limited sample, demonstrated an 80 percent reliability. The two case-patients with confirmed diagnoses of hepatitis A and mononucleosis may reflect on the specificity of this test. Based on the antibody response of all other case-patients, the rise in titer for patient No. 4 between days 88 and 94 after the reported onset of illness suggests the possibility of concomitant illness and a more recent exposure to *E. canis* antigen.

Clinical characteristics of case-patients were markedly similar to patients with RMSF and to controls. Oklahomans with symptoms of fever, headache, rash, and a history of tick-bite during the period April through August will prompt submission of samples for RMSF testing. Clinical characteristics among case-patients could be biased as a result of selection for RMSF and may not be representative of the disease as it exists in the general population. Differences such as the increased frequency of low white blood count in case-patients, when compared to controls, are likely to be valid. Rash occurred less frequently in case-patients than in controls, and when present was not of the typical pattern described for RMSF.

The results of this study implicate ticks as vectors for transmission of infection to case-patients. In spite of a selection bias for increased frequency of tick exposure in the population from which both case-patients and controls were drawn, a significantly higher proportion of case-patients reported finding ticks on their body than did controls. Previous reports indicate 50 percent to 100 percent of patients with serologic evidence of *E. canis* infection had tick exposure.[5–9] The recognized vector for *E. canis* in dogs in the United States, *R. sanguineus* (the brown dog tick), is reported to feed on humans infrequently suggesting that this species is of questionable importance in transmission of infection to humans. When the agent of human ehrlichiosis is isolated from a human patient, further studies of other tick vectors are indicated. If the agent is identified as *E. canis*, then the possibility of chronic infection of humans should be investigated, as chronic *E. canis* infection has been documented in dogs.[16]

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Self-Rated Health and Mortality in the NHANES-I Epidemiologic Follow-up Study

ELLEN L. IDLER, PHD, AND RONALD J. ANGEL, PHD

Abstract: The ability of self-rated health status to predict mortality was tested with data from the National Health and Nutrition Examination Survey (NHANES-I) Epidemiologic Follow-Up Study (NHEFS), conducted from 1971–84. The sample consists of adult NHANES-I respondents ages 25–74 years (N = 6,440) for whom data from a comprehensive physical examination at the initial interview and survival status at follow-up are available. Self-rated health consists of the response to the single item, "Would you say your health in general is excellent, very good, good, fair, or poor?" Proportional hazards analyses indicated that, net of its association with medical diagnoses given in the physical examination, demographic factors, and health related behaviors, self-rated health at Time 1 is associated with mortality over the 12-year follow-up period among middle-aged males, but not among elderly males or females of any age. (*Am J Public Health* 1990; 80:446–452.)

Introduction

A rather intriguing and consistent finding concerning the predictors of mortality has been reported by several researchers in recent years. Studies based on four different sets of data indicate that a simple, one item self-report of health status is as powerful a predictor of mortality as more detailed health status indicators.[1–3] In these studies the response to some variant of the question, "Would you say your health is excellent, very good, good, fair, or poor?" is significantly associated with the risk of mortality over periods of four to nine years, even when prior information about health and all other major known risk factors are accounted for. Without good baseline health indicators, this finding would not be particularly surprising. But the varieties of methods used and the consistency of the findings in these studies leads to the rather profound realization that what we learn from this simple question carries a great deal of clinically relevant information, making it as powerful a predictor of mortality as simultaneous reports of serious chronic disease or functional disability.

The first of these studies was published in the *American Journal of Public Health* in 1982.[1] In it, Mossey and Shapiro reported that in a large elderly Canadian sample of the province of Manitoba, respondents' self-ratings of health outweighed the extensive health care utilization and medical history data available in predicting survival over a seven-year period. The editor of the Journal, realizing the importance of the finding, highlighted it in an accompanying editorial.[4]

Subsequent studies in California, Connecticut, and Iowa have replicated the finding. Adults in Alameda County, California, who assessed their health as poor had increased relative risks of mortality of approximately 1.95 compared to those who said they were in excellent health, even when self-reports of chronic and acute illnesses, sex, age, income, education, health practices, and social networks were controlled.[2] Elderly adults in New Haven, Connecticut, and Iowa and Washington Counties, Iowa, had increasing risks of mortality for every level of self-assessed health other than excellent, despite controls for chronic conditions, disability, pain, medications, demographic factors, and health practices.[3] Together, these studies suggest that individuals have access to very important internal information with which they are able to make informed predictions of their probability of survival.

The primary methodological problem that makes attempts to assess the medical significance of self-reported health status difficult arises from the inability to unambiguously control for actual physical health status. In models in which self-assessments are used as predictors of some outcome, their impact may in fact be due to inadequately measured physical health status. In other words, it is difficult to establish that self-assessed health has an influence above and beyond that it shares with actual physical health. The studies mentioned above all took different approaches to isolating the net effect of self-assessed health, but all of these controls for physical health relied heavily on the self-reports of respondents and, as a result, none could be considered ideal controls for "objective" health status. Even the Canadian study, which depended to a smaller extent on respondent information, substituted records of medical care utilization which the authors acknowledged were subject to biases of their own.[1,4] No population-based study of self-assessed health and mortality has had as its physical health status control the results of a comprehensive, standardized physical examination carried out by trained physicians. Such an examination could be considered objective because it would consist of observations made by an external and impartial observer, and because a single standard would be applied to all participants in the study. Only this form of measurement would have the possibility of detecting untreated, undiagnosed, and otherwise unreported conditions.

Methods

Such a methodology is precisely the one employed by the National Center for Health Statistics's National Health and Nutrition Examination Survey (NHANES-I),[5] conducted from 1971 to 1975, and the NHANES-I Epidemiologic Follow-up Study (NHEFS),[6] conducted from 1982 to 1984. Specifically, the NHANES-I provides data from a detailed physical examination of a 20 percent subsample of the first-stage NHANES-I sample (N = 3,854), and an additional sample which is called the augmentation component (N = 3,059); together these make up a baseline sample of 6,913 adults ages 25 to 74 years who constitute a probability sample of the United States population. The first-stage NHANES-I sample deliberately oversampled certain population subgroups believed to be at risk for nutrition-related health

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problems: persons living in poverty areas, women of childbearing age, and the elderly. This sample is thus simultaneously very heterogeneous and nationally representative.

The follow-up to the NHANES-I, the NHEFS, successfully traced 92.9 percent of the original NHANES-I cohort;[7] for the subsample analyzed here the proportion traced was slightly higher, 93.2 percent. Those who were lost to follow-up tended to be younger, especially Black males and White females.[7] Respondents are included in the present analysis if they were successfully traced (by being contacted in person or having a proxy and/or a death certificate to verify their death), regardless of whether or not they completed the follow-up interview. This yields a usable sample size of 6,440. We might note that tracing over 90 percent of the sample 10 years after the initial survey is a truly commendable feat and it makes these data one of the best sources available for studying the predictors of mortality. The death rate for the subsample was 12.9 percent, somewhat lower than the 15.1 percent reported for the total sample.[7] This is likely due to the fact that the augmentation component was added to the sample at the end of the NHANES-I data collection period, thus making follow-up periods shorter for a substantial portion of our sample. The outcome variable used in this study is survival time following the initial interview.

The NHANES-I and the NHEFS are ideal for our purposes because they contain a fully objective source of information on the respondent's medical diagnoses at Time 1, as well as the respondents' self-ratings of health. The physicians made direct physical examinations of the respondents, including standing and sitting blood pressure and pulse, examination of the ears, head, eyes, mouth, neck, abdomen, major and minor joints, and skin, percussion of the liver, and auscultation of the heart. Based on positive responses to screening questions, the physician then administered supplemental questionnaires for arthritis, respiratory, and cardiovascular diseases. In addition, the physicians had immediately available to them the results of the following laboratory and other tests: hemoglobin, hematocrit, red and white blood cell counts, height and weight, body and skinfold measurements, audiometry, spirometry, electrocardiogram, single breath diffusing capacity, goniometry, and X-rays of the chest, hand and wrist, hips and knees.[5] The examination is as detailed as one could reasonably expect in a population survey of this scale. The physicians recorded their findings by coding any abnormalities they found in terms of the International Classification of Diseases Revision Eight (ICD-8) categories[8] and judged the severity of these conditions on a three-point scale as either minimum, moderate, or severe.[9] To create a set of variables which would summarize this information we grouped the ICD diagnosis codes by the ICD's own 15 broad categories: infectious diseases, cancers and neoplasms, metabolic disorders, blood diseases, mental disorders, diseases of the nervous, circulatory, respiratory, digestive, and genitourinary systems, diseases of the skin and musculoskeletal systems, congenital anomalies, symptoms of ill-defined conditions, and accidents.*

An individual received a score of one for that category if they had one or more diagnoses present; the score was then weighted by one, two, or three for the most severe grading the physician had given that individual for any diagnosis in that ICD category. Respondents were scored zero if they received no diagnosis in a category. In addition, to account for the additive effects of comorbidity, we created a variable that consists of the sum of unweighted separate diagnoses each respondent received. This complete set of variables is introduced first into all models as a control for medical diagnoses. It represents a distinct alternative to the methods used in previous studies in that it distills a large amount of information from a comprehensive physical examination that was conducted by a small team of physicians in a structured and standardized manner, and it incorporates a weight for the clinical severity of the condition, a feature which even the most accurate or extensive self-report data cannot claim.

Certainly a physical examination of the sort administered in the NHANES-I is not unerring. In a single cursory examination, the physician is likely to miss subtle conditions or syndromes in their prodromal state. Diagnosis is a complex task that depends on the physician's judgment and any single assessment of a person's state of health is always subject to observer variability.[10] The examination represents no more than the physician's informed clinical judgment arrived at after a brief encounter with a new patient. However, because the examination was performed by an external, independent, medically trained observer, and because these observers relied on standardized criteria for assessing the presence or absence of specific conditions, these diagnoses avoid confounding the impact of self-rated health on mortality with that of self-reported controls for health.

The independent variable of primary interest in this study is the response to the single item, "Would you say your health in general is excellent, very good, good, fair, or poor?" The question wording used here, and the labels for the five response categories differ slightly from those used in other studies;[1-3] in some cases the categories are excellent, good, fair, poor, and bad; sometimes only four categories are included and "poor" is omitted; sometimes the eliciting question contains a relativizing phrase such as "compared to other people your age . . ." The significance of these alternate forms in unknown; however, to date no two published studies have used the same form, although all have reported similar effects. Because of low frequencies in the "poor" category, the "poor" and "fair" categories were combined in this analysis, reducing the five response categories to four; the responses are treated as a set of n-1 dummy variables with the omitted category being "excellent" health.

The following sociodemographic variables are included in the analysis as controls: age (treated as a continuous variable); and dummy variables for race, 9 to 12 years of education or 13 or more years (compared with 8 or fewer), marital status, income less than $5000, income less than $15,000, and being currently employed. Health behavior variables included dummy variables for current smoking and past smoking (compared with never having smoked); an index of alcohol use based on the frequency of consumption multiplied by the number of drinks usually consumed, weighted by their alcohol content; the physician's indication of the presence of obesity; and a measure of inactivity based on frequency of exercise and other physical activity.

Proportional hazards models based on the SAS statistical procedure PHGLM[11,12] are used to estimate the effect of self-rated health on the probability of survival during the follow-up period. This procedure was used because of the censored nature of the data and because we wish to identify the predictors of survival time over the entire follow-up period. A complete, parallel set of logistic regression analyses on the dichotomous mortality outcome was also performed, which in effect measured the probability of survival at just one time point, the end of the follow-up period. The results of these logistic regressions were nearly identical to those from the survival analyses, hence we present only the latter. Odds ratios are calculated from the hazard models by taking the antilog of the estimated coefficients.[13] Three models are estimated, one containing just the physician's examination variables and self-rated health, one which adds sociodemographic variables in a block, and a third which adds health behavior variables. In this way, any changes in the relation between self-assessed health and mortality can be attributed to the various control variables.

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*Two additional categories, for pregnancy complications, and perinatal morbidity and mortality, were omitted as irrelevant and of negligible frequency.

Results

Figures 1 and 2 show the unadjusted relationship between self-assessed health at the time of the first interview and the probability of survival over the 12-year follow-up period for men and women. These figures show that: 1) mortality risks increase for individuals with progressively poorer self-perceptions of health; and 2) the risks are apparently greater for males than for females. Only 60 percent of males with poor self-assessed health survive to the end of the 12-year period, while nearly 80 percent of the women do. Further, the lines are more straight than curved indicating that the increased risk of mortality is spread throughout the follow-up period. Because of the differences in the risk of mortality, all analyses are performed separately by sex.

The full proportional hazards models are presented in Tables 1 and 2. In the first models, which contain only physician examination variables as controls, self-assessed health explains a significant amount of the remaining unexplained variation in survival time; odds ratios for the extreme category are 3.3 for men and 2.1 for women. There are also progressively elevated risks for the intermediate categories for men, but not for women.

The models which add sociodemographic variables reveal a diminished effect of self-assessed health on mortality for men, with lower odds ratios of 1.7 for poor/fair health and 1.5 for good health. Among the women all effects of self-rated health disappear. Stepwise logistic regressions not shown here reveal that age accounts for most of the association between self-assessed health and mortality for men and for all of the association for women. In this sample older people were more at risk for both poor self-assessments of health and mortality. Among the males, some of the apparent initial effects of self-assessed health on mortality were also due to confounding effects of employment, marital status, and race. Taking all these variables into account reduces the odds ratio for the extreme category by half.

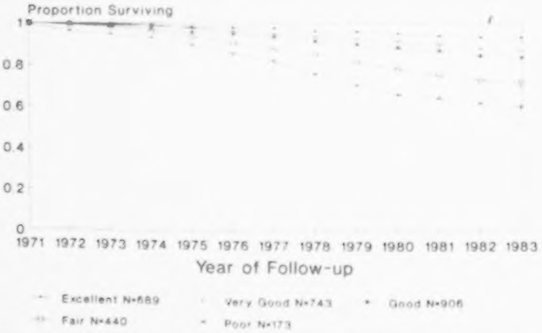

FIGURE 1—Male Mortality in NHANES-I Epidemiologic Follow-up Study by Levels of Self-rated Health

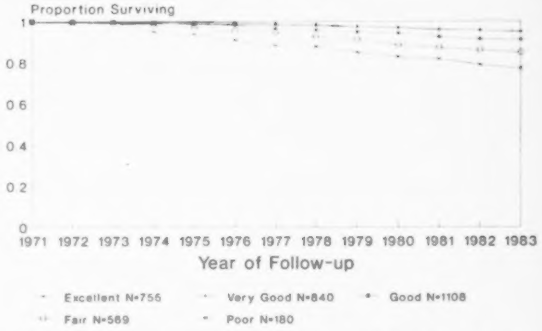

FIGURE 2—Female Mortality in NHANES-I Epidemiologic Follow-up Study by Levels of Self-rated Health

It should be noted that the model with sociodemographic factors taken into account fits the data much better than the model with health variables alone. The models' likelihood ratio chi squares (G^2) reported at the bottom of Tables 1 and 2 double when these variables are added, and the difference in G^2 is highly significant (588.2, 27 df - 285.9, 19 df = 302.3, 8 df, $p < .000$ for men; 429.0 27 df - 211.8, 19 df = 217.2, 8 df, $p < .000$ for women).

The third model introduces variables for smoking, alcohol use, obesity, and lack of exercise. The inclusion of these factors does not alter the already nonexistent relationship for women. It further reduces the association for men, so that only the extreme category contrast remains. A second stepwise procedure (not presented) revealed that the addition of current smoking to the model had the biggest impact on the association; males who smoke were both more likely to give poor self-assessments of health and to die. The final adjusted odds ratio for the poor/fair category in the men's hazard model is 1.5 (95% CI = 1.1, 2.2).

These models too have a substantially better fit than the sociodemographic models which preceded them. Adding the measures of health behaviors adds 54.1, 5 df ($p < .000$) to the G^2 for men, and 12.3, 5 df ($p < .05$) for women.

Given the literature on the ways in which self-ratings of health vary according to the respondent's social location, certain interactions were considered. We have already seen that self-rated health predicts mortality marginally for men, but not at all for women. But does it predict mortality better for some men (or women) than others? The issue of age is especially salient given that so many of the previous studies have been done with elderly samples. To test an age-self-rated health interaction, we divided the males and females into three groups, those ages 25 to 44, 45 to 64, and 65 to 74. Because deaths were extremely rare (31 males and 27 females) in the under-45 age group, making parameter estimates from these models unreliable, we do not report the results from the youngest group.

The results can be seen in Tables 1 and 2. A substantial difference emerges between the middle-aged and elderly men. The relation between self-rated health and mortality for the male sample as a whole disappears altogether for elderly men, and emerges very strongly for middle-aged men. In the

TABLE 1—Proportional Hazards General Linear Model Coefficients for Survival Time in the NHANES-I Epidemiologic Follow-Up Study (NHEFS) by Levels of Self-Rated Health, Males

	Model 1 Physician Controls	Model 2 Sociodemographics	Model 3 Health Behaviors	Model 3 By Age Group	
				45–64	65–74
ICD-9 Categories					
Infection	−.331 (.220)	−.214 (.222)	−.218 (.216)	−.554 (.418)	−.109 (.304)
Neoplasm	.249 (.105)	.108 (.114)	.125 (.115)	.228 (.213)	.131 (.154)
Endocrine	−.208 (.097)	−.111 (.100)	−.155 (.111)	.012 (.149)	−.309 (.182)
Blood	.050 (.417)	−.055 (.546)	−.141 (.572)	−6.160 (27.3)	.145 (.758)
Mental	−.024 (.165)	.047 (.159)	−.064 (.159)	−.072 (.206)	−.117 (.280)
Nervous	.007 (.072)	−.038 (.073)	−.004 (.073)	.070 (.114)	−.120 (.103)
Circulatory	.242 (.054)	.120 (.059)	.137 (.059)	.202 (.091)	.058 (.086)
Respiratory	.154 (.069)	.190 (.068)	.167 (.070)	.134 (.124)	.201 (.094)
Digestive	.170 (.098)	.112 (.104)	.144 (.109)	.142 (.160)	.084 (.155)
Genital	−.852 (.730)	−.574 (.692)	−.431 (.681)	.084 (.699)	−1.859 (10.2)
Skin	−.050 (.106)	−.122 (.117)	−.161 (.119)	.036 (.185)	−.278 (.163)
Muscle	−.011 (.061)	−.119 (.064)	−.114 (.064)	−.038 (.101)	−.161 (.090)
Congenital	−.249 (.188)	−.280 (.206)	−.369 (.224)	−.822 (.586)	−.334 (.275)
Illness Symptoms	.207 (.082)	.278 (.087)	.260 (.088)	.326 (.127)	.184 (.140)
Accident	−.270 (.136)	−.156 (.131)	−.225 (.132)	.081 (.167)	−.916 (.444)
Sum of conditions	.147 (.042)	.089 (.047)	.068 (.047)	−.046 (.075)	.159 (.066)
Self-rated Health					
Poor/Fair	1.203 (.167)	.552 (.184)	.436 (.189)	1.026 (.330)	−.071 (.258)
Good	.748 (.166)	.396 (.178)	.292 (.182)	.793 (.322)	−.144 (.247)
Very good	.543 (.176)	.287 (.187)	.251 (.190)	.646 (.342)	−.248 (.259)
Sociodemographics					
Age		.059 (.005)	.066 (.006)	.060 (.014)	.083 (.025)
White		−.327 (.127)	−.309 (.129)	−.487 (.195)	−.215 (.191)
High school education		.136 (.115)	.086 (.116)	.077 (.176)	.044 (.176)
College education		.194 (.158)	.158 (.160)	.131 (.241)	.405 (.236)
Married		−.387 (.116)	−.305 (.118)	−.409 (.184)	−.083 (.178)
Income < $5000		.451 (.180)	.384 (.181)	.464 (.269)	.260 (.307)
Income < $15000		.273 (.157)	.210 (.157)	.255 (.213)	.076 (.295)
Working		−.460 (.117)	−.460 (.118)	−.370 (.181)	−.444 (.180)
Health Behaviors					
Present smoker			.616 (.135)	.835 (.230)	.551 (.191)
Past smoker			.027 (.141)	.426 (.244)	−.222 (.187)
Alcohol use			.001 (.000)	.001 (.000)	−.000 (.000)
Obesity			.209 (.124)	.156 (.186)	.220 (.183)
Inactivity			.090 (.044)	.139 (.069)	.090 (.065)
N	2931	2796	2772	1225	514
G^2	285.9	588.2	642.3	186.6	95.3
degrees of freedom	19	27	32	32	32

latter group, odds ratios for the poor/fair, good, and very good categories are 2.8 (95% CI = 1.5, 5.3), 2.2 (95% CI = 1.2, 4.1), and 1.9 (95% CI = 1.0, 3.7), respectively. Self-rated health does appear to predict mortality very strongly for middle-aged men and it predicts better for them than for any other group.

Because we were interested in the relative contributions of the self-rated health and physician examination variables to the explained variance in mortality, we compared the fit of the full models (Model 3 in Tables 1 and 2) with the fit of models from which we had alternately removed the variables for self-assessed health and physician-assessed health. This would allow us to assess the relative contribution of the two health status measures by their contribution to the G^2 for the full model. We estimated these models for the whole sample, and also by age group. Our findings varied, both by sex and by age, and were somewhat surprising: for the women in the sample, neither the physicians' diagnoses nor self-assessed health make any significant contribution to the model G^2 when sociodemographic and health behavior variables are controlled. While some of the individual physician measures were significant predictors of female mortality (notably diagnoses of circulatory and mental disorders), the removal of the entire set of diagnoses from the model failed to reduce the G^2 by a significant amount. For the males, on the other hand, the physician measures (diagnoses of circulatory and respiratory disorders, and ill-defined illness symptoms in particular) did contribute to G^2, both for the male sample as a whole (642.3, 32 df - 606.2, 16 df = 36.3, 16 df, $p < .01$) and for the elderly men (95.3, 32 df - 57.5, 16 df = 37.8, 16 df, $p < .01$). In none of these groups did the block of self-assessed health variables have an impact on G^2. Among the middle-aged men, however, where the effect of self-assessed health on mortality was strong, self-assessed health does alter G^2 (186.6, 32 df - 174.8, 29 df = 11.8, 3 df, $p < .01$). The effect of the physician-assessed measures within this group is marginal (186.6, 32 df - 162.0, 16 df = 24.6, 16 df, $.10 > p > .05$). These comparisons corroborate the findings regarding the importance of self-assessments of health among middle-aged men, and show the findings from the physicians' examinations to be better predictors of male mortality than female mortality.

To focus a little more closely on the role of self-assessed health vis-a-vis the other important predictors of mortality, we compared the odds ratios and partial Rs for the various levels of self-assessed health with those for other variables in

TABLE 2—Proportional Hazards General Linear Model Coefficients for Survival Time in the NHANES-I Epidemiologic Follow-Up Study (NHEFS) by Levels of Self-Rated Health, Females

	Model 1 Physician Controls	Model 2 Sociodemographics	Model 3 Health Behaviors	Model 3 By Age Group	
				45–64	65–74
ICD-9 Categories					
Infection	−.198 (.223)	.209 (.223)	.163 (.223)	.544 (.274)	.071 (.553)
Neoplasm	.171 (.203)	.023 (.211)	.037 (.210)	.038 (.538)	−.069 (.272)
Endocrine	.081 (.079)	.109 (.080)	.142 (.085)	.047 (.149)	.201 (.113)
Blood	.022 (.553)	−.039 (.534)	−.021 (.531)	−7.601 (86.3)	.171 (.483)
Mental	.189 (.183)	.440 (.183)	.419 (.187)	.038 (.372)	.861 (.281)
Nervous	.042 (.102)	−.036 (.105)	−.045 (.105)	.186 (.183)	−.146 (.137)
Circulatory	.443 (.067)	.241 (.071)	.253 (.072)	.360 (.131)	.235 (.093)
Respiratory	−.144 (.148)	.010 (.141)	−.034 (.141)	.100 (.242)	.002 (.181)
Digestive	.051 (.186)	.156 (.183)	.105 (.191)	.180 (.325)	.015 (.268)
Genital	−.197 (.353)	.008 (.352)	.028 (.354)	−6.614 (62.4)	.221 (.426)
Skin	.002 (.190)	.102 (.192)	.088 (.195)	.147 (.307)	.136 (.277)
Muscle	.268 (.071)	.012 (.073)	.020 (.073)	.074 (.145)	.029 (.089)
Congenital	−.401 (.305)	−.183 (.299)	−.236 (.293)	−6.458 (29.3)	.008 (.290)
Illness Symptoms	.112 (.104)	.000 (.106)	.012 (.107)	.064 (.204)	−.068 (.148)
Accident	−.164 (.327)	.056 (.315)	.051 (.315)	−.080 (.527)	.393 (.668)
Sum of conditions	.061 (.046)	−.009 (.049)	−.012 (.049)	−.096 (.105)	.021 (.060)
Self-rated Health					
Poor/Fair	.735 (.188)	.184 (.196)	.204 (.202)	.256 (.354)	.084 (.287)
Good	.324 (.187)	−.029 (.191)	−.010 (.194)	−.009 (.330)	−.037 (.278)
Very good	−.129 (.219)	−.321 (.222)	−.246 (.224)	−.195 (.372)	−.456 (.339)
Sociodemographics					
Age		.067 (.006)	.071 (.007)	.049 (.019)	.049 (.029)
White		−.458 (.151)	−.466 (.152)	−.485 (.259)	−.341 (.214)
High school education		−.022 (.138)	−.056 (.139)	.195 (.248)	−.097 (.185)
College education		−.020 (.191)	−.064 (.193)	−.051 (.377)	−.387 (.269)
Married		−.391 (.128)	−.378 (.129)	−.322 (.229)	−.304 (.170)
Income < $5000		.271 (.228)	.301 (.230)	.553 (.356)	.085 (.395)
Income < $15000		.114 (.213)	.123 (.214)	.046 (.328)	.202 (.387)
Working		−.380 (.164)	−.381 (.165)	.053 (.227)	−.813 (.400)
Health Behaviors					
Present smoker			.384 (.145)	.596 (.222)	.455 (.225)
Past smoker			.332 (.172)	.206 (.330)	.335 (.227)
Alcohol use			.000 (.000)	.000 (.000)	.000 (.001)
Obesity			−.175 (.132)	.158 (.234)	−.305 (.176)
Inactivity			.040 (.057)	.059 (.104)	.015 (.074)
N	3417	3267	3242	1358	557
G^2	211.8	429.0	441.3	74.2	61.4
degrees of freedom	19	27	32	32	32

the full models. For the male sample as a whole, age is the most important predictor of mortality, followed (in this order) by smoking, unemployment, alcohol use, illness symptoms, marital status, race, respiratory diagnoses, poor self-assessed health, circulatory diagnoses, low income, and lack of exercise. Among the middle-aged men, poor self-assessed health ranks fourth as a predictor, behind (in this order) alcohol use, age, and smoking. Among the women the order is age, circulatory diagnoses, race, marital status, smoking, unemployment, and mental disorders. We conclude that health behaviors (especially smoking and high levels of alcohol use) and demographic variables (particularly employment) are more important predictors of mortality in age-adjusted models than either physician- or self-assessed health. Among the physician measures, however, we would note that ill-defined illness symptoms, and diagnoses of respiratory and circulatory problems are the best predictors of mortality for men, and diagnoses of circulatory and mental disorders are most important for women.

In understanding the impact of self-rated health on mortality, one of the other concerns has been that poor self-assessments of health may simply be reflecting the greater severity level or better respondent knowledge of conditions which are leading to imminent mortality.[1] If people with poor self-assessed health only survive for a short follow-up period then the effects of self-assessed health are more easily explained as inadequate controls for physical health status than as indicators of some more subtle process.

To test this possibility, we identified deaths which occurred in three "early mortality" periods of two, three, and five years following the year in which the respondent entered the NHANES-I study, and compared them with the deaths that occurred in the remaining years. We then performed logistic regression analyses with the full set of variables on just the subset of deaths to see if self-assessments of health predicted early mortality better than late mortality. The results (not shown) revealed that, among both males and females, self-rated health did not predict early mortality better than late. This confirms the findings of Mossey and Shapiro,[1] who concluded that ". . . while objective health status may change over time, self-ratings of health represent a relatively stable perception held by the individual." (p 804) This suggests that the effects of self-assessed health on mortality cannot be explained by superior respondent knowledge of the poor prognoses for existing conditions that are leading to imminent mortality.

Discussion

These analyses have shown that the ability of self-rated health to predict mortality in this national sample is in large part explained by the sociodemographic characteristics, health risk behaviors, and the medical diagnoses of the respondents. Only among middle-aged men was there any substantial independent increased risk of death from self-assessments of health after these other factors were taken into account. In this NHANES-I sample, older people, smokers, alcohol abusers, the unmarried, the unemployed, and Blacks were initially less likely to give positive assessments of their health and subsequently less likely to survive, and these factors, by and large, account for the initial association of self-rated health with mortality.

Our findings, then, do not fully corroborate the findings of the earlier studies reviewed above in which self-assessments remained strongly significant predictors of mortality in full samples, in spite of controls for physical health indicators and other factors. Our analysis differs from these earlier studies in two major respects: 1) the NHANES-I provides an alternate indicator of physical health, medical diagnoses; and 2) the analysis is based on a large and heterogeneous national sample. The previous studies were of localized and/or elderly samples, and employed controls for physical health status that were at least in part based on self-reports. We attribute the differences between our results and those of previous studies to these differences in study design.

No other population-based prospective study of self-assessed health has had medical diagnoses derived from a standardized medical examination as statistical controls for physical health status. The NHANES-I detailed examination could potentially yield information on previously undiagnosed as well as treated and self-reported illness. Indeed, this was a primary purpose of the NHANES-I: ". . . to obtain information on unrecognized and undiagnosed conditions—in some cases, even nonsymptomatic conditions."[5] (p 3) In studies which depend on self-reports of conditions for their measure of physical health status, the same optimistic or pessimistic response bias may influence specific reports of conditions as well as global assessments. The current study is unique in basing its controls for physical status on information gathered from an external observer and hence applies the same standard to all respondents, regardless of their race, income status, education level, or access to medical care.

Having said this, it should be noted that however theoretically advantageous these controls for medical diagnoses are, they do not on their own eliminate the association of self-assessed health with mortality in this sample. In fact, they explain very little of the observed variation in survival. Perhaps this should not be surprising; this was a fundamentally healthy young and middle-aged sample and few cases of serious illness were diagnosed. Moreover, in making their diagnoses, the physicians were not asked to consider longevity as a criteria. In any case, among both men and women, self-assessed health remained a strong predictor of survival when the diagnoses alone were controlled. In this analysis, the relationship of self-assessed health to mortality was better explained by age (primarily) and, to a lesser extent, by smoking, alcohol abuse, unemployment, race, and marital status than it was by its confounding with physical health status as judged by these physician examinations.

The second major difference between the present study and those which have preceded it concerns the sample. The NHANES-I sample is large and heterogeneous, quite unlike the community, county, or even provincial samples discussed earlier. One concrete example of the differences is in the role that age plays in the analyses. In the Connecticut and Iowa samples,[3] which contained only elderly respondents, age acted as a suppressor variable on the relation between self-assessed health and mortality; that is, because older (elderly) respondents gave better self-assessments of health at any given level of objective health status than younger (elderly) respondents, controlling for age in the analysis *increased* rather than decreased the effect of self-assessments on mortality. In the NHANES-I sample of young and middle-aged adults and only relatively young elderly, the effect of age is very much the opposite: adding a main effect of age to the models sharply curtails or eliminates the association of self-assessed health with mortality.

Only in one age/sex group do we see the hypothesized effect. In a sense, however, this is the group for which the public health significance of the findings is potentially greatest, since male mortality in this age group greatly exceeds female mortality, and since most deaths which occurred in this group could be considered premature and conceivably preventable. The limited findings in this study may provide some clue to the origins and meaning of these subjective perceptions. One might speculate that the apparently accurate predictions of survival made by the middle-aged male respondents in this study could be derived in part from a knowledge of their family's history. Having had a close family member who suffered a premature death from heart disease or cancer could lend ominous significance to symptoms such as a persistent cough or a mildly high cholesterol level which might appear unimportant to someone from a long-lived family. One might also observe that the strength of the association in this age group underscores the fact that, in our society, "excellent" health is the norm for the non-elderly, and any deviation from the norm indicates a potentially serious risk.

Even in comparison with the urban and nonelderly Alameda County sample with which it has more in common, however, the NHANES-I sample is diffuse and geographically unbounded. It encompasses the complex ethnic and cultural diversity of the nation. The cognitive processes by which individuals make global judgments about internal states are by no means well understood, but there is strong evidence that they somewhat consistently vary from one social, ethnic, or cultural setting to another.[14–17] Self-assessments appear to be made by some process of reference-group comparison, in which the individual chooses an appropriate group and judges the state of his or her own health by the health levels typical of the group in question. In other words, these judgments have no absolute reference point. The national sample in a sense removes these judgments from the cultural and community contexts in which they were formed, causing the variation in mortality to be explained primarily by the more objective and also more diverse physical health, demographic and health behavior factors. This analysis should remind us that demographic and health risk behavior factors remain potent predictors of mortality.

Self-assessments of health should continue to be of interest as mortality predictors, not only within homogeneous community samples, but also in large national samples where their impact appears more limited. Self-ratings of health are inexpensive, easy to obtain, and already available in nearly all health surveys. Even if they act as sensitive indicators of mortality risk only for certain subgroups in the population, they should remain valuable tools for health researchers.

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Lowell Levin to Edit New International Journal

Lowell S. Levin, EdD, MPH, professor of public health at the Yale University School of Medicine, has been named editor of the new *International Journal of Iatrogenic Complications*. The quarterly journal, associated with the Copenhagen-based International Society for the Prevention of Iatrogenic Diseases, will begin publishing in February 1991.

Dr. Levin explains that although "iatrogenic" translates literally as "physician-produced," the study of such problems has come to emcompass the negative consequences brought on by any segment of the increasingly complex health care system. When a part of that system harms a patient either by accident, negligence or imcompetence, the ramifications go beyond the medical, to the legal, social and economic. Hence, each article submitted to the new periodical will be reviewed by a jury of three readers from various disciplines, rather than the two readers used by many journals. In addition to physicians, the reviewer roster will include attorneys, economists, social scientists, epidemiologists and others, drawn from a broad geopolitical spectrum.

Dr. Levin is an internationally recognized authority in the field of self-care, which emphasizes education to empower the public in health matters and to prevent illness. He has helped pioneer this concept in both the developed and developing worlds, and has served on various World Health Organization (WHO) committees, including the WHO Committee on Information and Education for Health.

Scholars or other researchers who would like to submit articles to the *International Journal of Iatrogenic Complications* may contact Dr. Levin at (203) 785-2849.

The Economic Impact of Injuries: A Major Source of Medical Costs

LINDA C. HARLAN, PHD, WILLIAM R. HARLAN, MD, AND P. ELLEN PARSONS, PHD

Abstract: Data from the 1980 National Medical Care Utilization and Expenditure Survey were analyzed to place the costs for injuries in the context of all medical costs and to describe the distribution by demographic and diagnostic groups. For the non-institutionalized population, injuries, which include intentional and unintentional, were the second leading cause of direct medical costs, accounting for $16,745 million in medical care expenditures and a major contributor to work loss and disability in the US. For the working-age population (17–64 years) injuries were the leading cost category ($11,341 million) and the third most costly category for persons 65 years of age and over ($3,479 million). The preponderance of costs were attributable to hospital-based care. Direct medical costs were disproportionately greater for males, White and other persons, and for those with household incomes less than $5,000. Injury morbidity also accounts for major indirect costs. Fractures accounted for the highest direct medical costs, greatest per capita charges (based on those with charges), and largest number of restricted activity days. These national estimates document the economic importance of injuries and direct public attention to policy imperatives related to research and prevention. (*Am J Public Health* 1990; 80:453–459.)

Introduction

Injuries are the most common cause of death in the United States for persons under the age of 44.[1] In 1986, unintentional injuries ranked as the leading cause of years of potential life lost before age 65.[2] This consideration and the high incidence of injuries and consequent disability have directed public attention to this problem.[1,3] Despite the frequency of accidental injury, there is a relative paucity of national data on medical costs and morbidity. There appears to be a major disparity between health costs (both direct and indirect) and research investment for prevention and management.[1] However, valid estimates of the national economic impact have not been available until recently[4,5] and these total estimates do not provide a full understanding of the impact of accidents and injuries in various age-sex groups. Information on direct medical costs and indirect costs from disability could provide an economic perspective that would buttress recommendations for greater public investment in research in this field.[1] Moreover, the distribution of costs across demographic and diagnostic categories affords a background understanding of utilization that could lead to focused efforts at prevention and cost containment.

To provide a picture of the national expenditures for injuries, we have analyzed data from the 1980 National Medical Care Utilization and Expenditure Survey (NMCUES). This survey provides the most recent national data on the costs of injuries, which include both intentional and unintentional injuries.

Methods

NMCUES, national survey of health expenditures and use of personal health services, was conducted by the National Center for Health Statistics and the Health Care Financing Administration during 1980 on a national sample of the civilian, non-institutionalized population of the United States. The survey included persons living in 6,600 households. The sample design was a multi-stage probability sample drawn from 135 primary sampling units representing the 50 states and the District of Columbia. Data on 17,123 persons were collected from February 1980 through April 1981. Data were analyzed for Whites and others in one racial category and Blacks in another. The "other" category included American Indians, Alaskan Natives, Asians, and Pacific Islanders. These individuals accounted for approximately 2.5 percent of the "White and other" category. The effect of "others" in the "White and other" category would be expected to be very small. Approximately 7 percent of the NMCUES sample reported themselves as Hispanic. The sample weight was constructed based on race, not ethnicity. The Hispanics are distributed across the racial categories, most being in the White racial category. The estimates obtained if Hispanics were analyzed separately would be non-representative.

Interviews were conducted with each household at approximately three-month intervals. The first two interviews were conducted during a personal household visit by an interviewer; the next two were conducted by telephone; and the final interview was a personal household visit. At the first interview, the respondent was given a diary and asked to record occurrences, conditions, and functional impairments. In addition, the respondent was to record all charges for illnesses and injuries irrespective of who paid these charges, i.e., Medicare, Medicaid, private insurance, or out-of-pocket expenses. Subsequent telephone and personal interviews focused on these diaries. The final response rate was 87.9 percent when attrition over the course of interviewing was taken into account. Detailed information on the survey design, sampling bias, instruments, and statistical analyses are available in several monographs.[5–8]

The public use files of NMCUES were used for these analyses after edits were performed as outlined in the NCHS methodological monograph.[7] Conditions were coded using the International Classification of Diseases, ninth revision, ICD-9[9] and special NCHS codes[10] and grouped into 18 broad diagnostic categories. This report focuses on the "injuries and poisonings" category. In this report, these are referred to as "injuries." The injury category comprised the diagnostic groupings and the relevant ICD codes shown in the Appendix.[9] Both intentional and unintentional injuries are included in these analyses. It was not possible to determine from the data the cause of the injury.

Direct costs for these analyses were defined as charges incurred and reported for inpatient and outpatient care,

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physician and allied health services, prescribed medication and selected medical equipment and other services. The reports of costs were not verified so it is likely that there was an underreporting of actual charges. Some groups probably underreported their charges more than others. Individuals covered by Medicaid, General Assistance or Veterans Administration benefits may not be informed by the payer of the actual charges and may thereby underestimate charges more than other groups.[5] In addition, these data include information only on civilian, non-institutionalized individuals. Costs related to individuals in hospitals, nursing homes, and other institutions or the military would not be represented here.

Disability attributable to these diagnostic groups was categorized into restricted activity (usual activity restricted by illness), bed disability (confined to bed more than half a day), and work loss (lost work because of illness). Restricted-activity days is the most inclusive category and include bed-disability days. Bed-disability days include all hospital days, even if a bed was not occupied. The complex sample design, sampling and respondent biases, and entry or attrition from the survey during the survey year were taken into account in the analysis using procedures described in a methodological monograph.[7] Use of these procedures permits national estimates to be made.

Results

Direct Costs

Direct medical costs for injuries comprised the second largest source of expenditures for medical care in the United States during 1980 (Figure 1). The estimated $16.8 billion for injuries represents 12 percent of all direct costs. This amount was exceeded only by an estimated $20 billion in charges for diseases of the circulatory system.[3]

The distribution of medical costs for the four leading condition categories is presented by age groups in Figure 2. Injury and poisoning ranked among the top cost categories for three of the four age groups (<17 years, 17 to 44 years, and >65 years). Injury costs were the greatest ($9 billion) for young adults (17 to 44 years). This diagnostic category was the third largest contributor to medical costs for older Americans (65 years and over) and was exceeded only by costs related to circulatory conditions and neoplasms. When only the working-age population (17–64 years) was considered, full-time or part-time workers had injuries and poisonings as their major source of medical costs. In contrast, persons not employed had greater charges from other diagnostic categories, generally chronic diseases.[6]

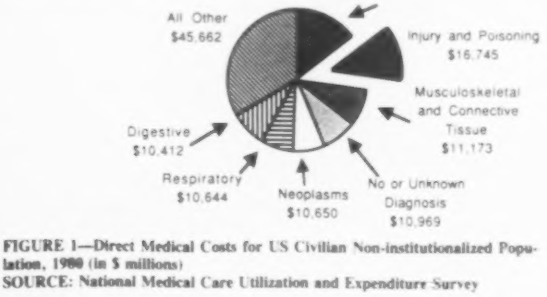


FIGURE 1—Direct Medical Costs for US Civilian Non-institutionalized Population, 1980 (in $ millions)
SOURCE: National Medical Care Utilization and Expenditure Survey

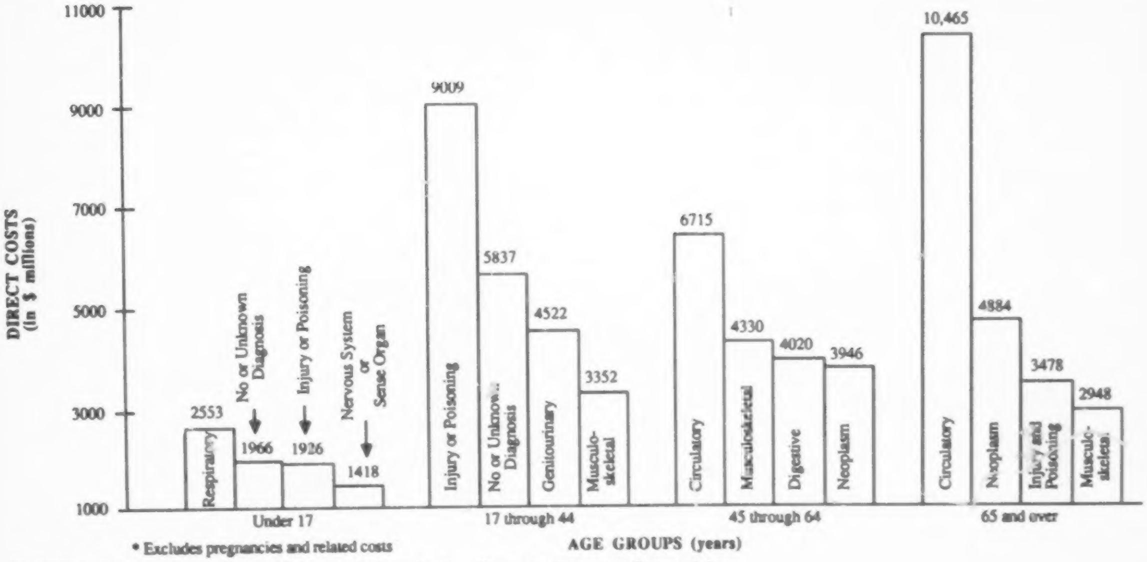
The proportion of medical costs attributable to injuries are presented by demographic characteristics in Figure 3. Medical costs for injuries are presented as the proportion of total direct medical costs for each referent group; the denominator is the total direct cost for all conditions for that group. Injuries accounted for a greater proportion of medical costs for males than for females. White and other persons also had proportionately greater medical costs attributable to injuries than did Black persons. The distribution of injury costs across income groups indicates that persons with family incomes less than $5,000 had a greater proportion of medical costs attributable to injuries than did those in the other income groups. The lower income group (<$5,000) comprised 7.3 percent of the total population, but their total injury costs accounted for 10.3 percent of all injury costs. Thus, the lowest income group had disproportionately greater costs for injuries and disproportionately higher health care expenditures.

The injury category was separated into 11 diagnostic groups which followed the ICD-9 diagnostic categories. The direct costs were determined for each of these groups. The costs attributable to each diagnostic group are presented in Table 1, as are the distributions by age group, sex, and race. Fractures, which includes skull fractures, were responsible for the greatest expenditures and generated $6.4 billion in charges or almost 40 percent of total costs for injuries. The early treatment of trauma ($2.6 billion) and management of the late effects of trauma ($1.7 billion) were also major sources of costs. Sprains and dislocations, wounds (including blood vessel injuries) and complications of medical and surgical care were the only other diagnostic groups leading to over $1 billion per year.

There were important variations in diagnostic group costs by age, sex, and race (Table 1). The costs attributable to fractures, the most costly category, varied by age group from 27 percent (<17 years) to 43 percent (>65 years). Fractures accounted for more costs for females (42 percent) than for males (35 percent). Black persons had 44 percent of injury costs attributable to fractures in contrast to 38 percent for White and other persons. The costs associated with early treatment of trauma were highest for the 45 to 64 year old group, for females, and for Black persons. This grouping includes multiple traumatic injuries, but excludes specific traumatic diagnoses such as fractures and burns. Costs associated with later effects of trauma were greatest for those over age 65 years, for females, and for White and other persons. Sprains and dislocations generated relatively low charges for those over 65 and for females. Wounds accounted for 21 percent of costs for those under 17 years and wounds and fractures together accounted for almost half (48 percent) of costs in this age group. Poisoning costs were overwhelmingly associated with females and were highest in the 17 to 64 year old group. Costs for burns were higher for males than for females and for those 45 to 64 years.

The immediate cause of the injury (e.g., vehicular accident, fall, attempted suicide, etc.) and contributing causes (e.g., alcohol ingestion, speeding, etc.) were not recorded in NMCUES. Therefore, only indirect inferences can be made about causation. The distribution of injuries suggest a relatively greater cost impact of trauma, i.e., sprains, wounds and head injuries, from leisure and work activity on young males and greater cost impact on older women from fractures and late sequelae of trauma.

These aggregate direct costs do not disclose the costs on a per occurrence or per person basis or the allocation of costs



* Excludes pregnancies and related costs

FIGURE 2—Direct Medical Costs for All Persons by Four Leading Diagnostic Categories, Cost, and Age

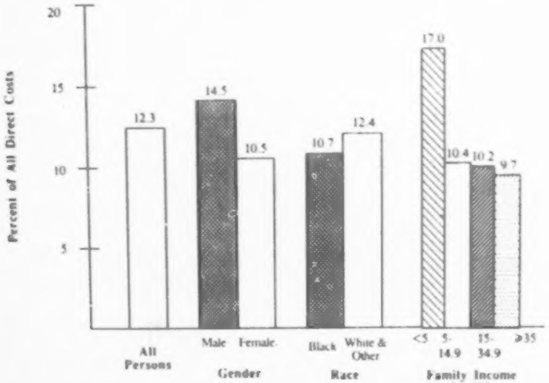


FIGURE 3—Proportion of Total Direct Medical Costs Attributable to Injuries by Demographic Groups
SOURCE: National Medical Care Utilization and Expenditure Survey

across components of care. Therefore, the per capita charges were determined for the diagnostic categories and divided into the charges for four types of health services: hospital care (including emergency room), ambulatory medical care (including outpatient physician charges), hospital physician charges, and medication charges (Table 2). The per capita charges are based on mean cost per individual having charges under each diagnostic and health service category. On a per capita basis, the highest charges were for management of the late effects of trauma and for fractures. For the other diagnostic groups, the rank order of per capita charges differed somewhat from the rank order of total direct charges (Table 1). Sprains had relatively low per capita costs but high aggregate costs because of greater frequency. On the other hand, burns had high per capita costs but low aggregate costs. The most striking finding regarding per capita charges is that the preponderance of charges for each diagnostic code (except sensory impairment) were generated by hospital care, which includes emergency room services. However, for the diagnostic categories of foreign bodies, burns, and sensory impairment, the estimates are unstable because fewer than 10 individuals in NMCUES were hospitalized in each category during the survey. The next highest category of charges were for physicians' hospital services. Only small charges ranging from $12 to $71 were reported for medications and from $57 to $224 for ambulatory care. Therefore, over 90 percent of per capita costs for most injury categories were related to hospital-based care.

Disability

Disability and work loss are indirect morbidity costs of medical conditions. Although total disability, work loss, and restricted-activity days were collected on each individual, a single day could be assigned to several conditions. Therefore, days lost for each condition were calculated by multiplying the weighted proportion of days lost due to the condition of interest by the total days lost. A national estimate of 932,300 person years of restricted activity and 292,300 person years of bed disability were attributable to injuries. These days represent about 10 percent of days attributable to all conditions during 1980 and exceed the estimates of disability for circulatory conditions for the same period.[11]

Restricted-activity and bed-disability days are presented for the diagnostic groups in Figure 4. Fractures were responsible for the greatest mean number of restricted-activity and bed-disability days at all ages. For persons age 65 years and older, the mean number of days approximately doubles. Similarly, two- to three-fold greater disability occurred for older adults with head and internal injuries, late effects of trauma, and complications of medical and surgical care. For sprains, burns, and early effects of trauma, there were no differences by age. On the other hand, poisonings, sensory impairment, and wounds had lower mean days of restricted activity or bed disability for older persons, but the frequency

TABLE 1—Direct Medical Costs for Diagnostic Codes by Age Group, Gender and Race: National Medical Care Utilization and Expenditure Survey, 1980

	Total Charges* (in $millions)		Age Group				Gender		Race	
Diagnostic Category	$	(%)	<17	17–44	45–64	65+	Male	Female	White & Other	Black
Fractures	$ 6421	(38.3)	524	3678	709	1510	3071	3350	5763	658
Early Complications of Trauma	2602	(15.5)	280	1171	510	641	1103	1499	2317	285
Late Effects of Trauma	1694	(10.1)	222	730	103	638	695	998	1668	26
Sprains, Dislocations	1599	(9.5)	240	944	295	120	1048	551	1438	161
Wounds (including blood vessels)	1565	(9.3)	406	849	150	160	1089	476	1456	109
Iatrogenic†	1152	(6.9)	87	586	185	295	634	519	1104	49
Head Injuries	816	(4.9)	53	680	55	29	633	183	692	124
Poisonings	383	(2.3)	30	226	80	47	26	357	348	35
Burns	281	(1.7)	36	51	190	4	225	55	264	17
Foreign Body Entry	132	(0.8)	26	54	30	23	70	62	107	25
Impairment of Senses	100	(0.6)	22	40	25	12	80	19	96	4
Total**	16745	(99.9)	1926	9009	2332	3479	8674	8069	15253	1493
(proportion of Total Costs)			(11.5)	(53.8)	(13.9)	(20.8)	(51.8)	(48.2)	(91.1)	(8.9)

*Does not include charges for dental services.
**Do not necessarily add to $16745 due to rounding.
†Complications of medical and surgical care.

TABLE 2—Per Capita Medical Charge for Individuals with Charges by Diagnostic Codes, National Medical Care Utilization and Expenditure Survey, 1980

Diagnostic Category	Hospital	Ambulatory	Hospital Physician	Prescription & Other Service
Fractures	4548	194	986	39
Late Effects of Trauma	3818	104	1005	33
Early Complications of Trauma	2198	87	112	23
Iatrogenic†	2138	113	1017	27
Burns	2104	67	985	18
Poisonings	1763	57	196	12
Wounds (including blood vessels)	1279	90	776	16
Sprains, Dislocations	1271	102	483	18
Head Injuries	1260	224	449	53
Foreign Body Entry	669	74	263	17
Impairment of Senses	205	175	1068	71
All Injuries	2879	123	817	27

†Complications of medical and surgical care

was low and the estimates unstable. In general, bed-disability days comprise less than half of restricted activity except for those over age 65 years who had fractures, head and internal injuries, and late effects of trauma. Although not reported here, the distributions of disability by gender and race were similar to the distributions of direct and per capita costs.

Work-loss days were estimated for the working age-population (17 to 64 years), both men and women. The total work loss attributable to injuries was 247,900 person years. This represents 14 percent of all work-loss days attributable to illness. The mean work-loss days by diagnostic groups are plotted in Figure 5. Fractures were responsible for the highest mean work-loss days. Sensory impairment, head and internal injuries, sprains, burns, and the early and late effects of trauma are the other injuries that resulted in a mean work loss of five days or more. There were important differences between men and women for several of the condition groups. Men had more work-loss days than women from all conditions except sensory impairment and fractures. Although not shown, there were no important trends by race, except for greater mean work-loss days for Black persons attributable to early and late treatment of trauma.

Discussion

These analyses confirm that injuries are a major source of medical costs, both direct medical charges and indirect costs of illness, as measured by disability days. Injuries are the second leading cause for direct medical costs in the civilian, non-institutionalized population. This cost burden is spread across all ages. Injuries are the leading contributor to medical costs for persons 17 to 44 years of age. Importantly, injuries also account for a considerable proportion of medical costs for Americans, age 65 years and over. For this older group, only circulatory conditions and neoplasms are responsible for higher costs. The general perception that medical expenses in older, non-institutionalized persons result primarily from chronic diseases and infections contrasts with these national findings.

These survey data underestimate the actual economic impact on the elderly because NMCUES did not survey institutionalized persons who represent an important proportion (about 10 percent) of those over age 65 years. As such, NMCUES does not address the cost burden related to institutionalized individuals. Hodgson and Kopstein have reported that the costs for this group are different than those living in the community.[4] Using a variety of data sources, they estimated that per capita expenditures for the elderly institutionalized and non-institutionalized population rank injuries as the 4th and 8th, respectively for women and men. These authors also found that although only 11 percent of the population is over the age of 65, they consume about 31 percent of the personal health care dollars. Therefore, the total cost estimates for the US population would be greater than estimated from this survey. Rice, *et al*, reported that when institutionalized and non-institutionalized elderly are included that diseases of the circulatory system account for 31 percent of the total direct costs.[12] Because the elderly contribute disproportionately to total costs and because chronic disease is responsible for a greater proportion of the cost when institutionalized persons are included, injuries would account for a somewhat lower proportion of medical costs for the elderly.

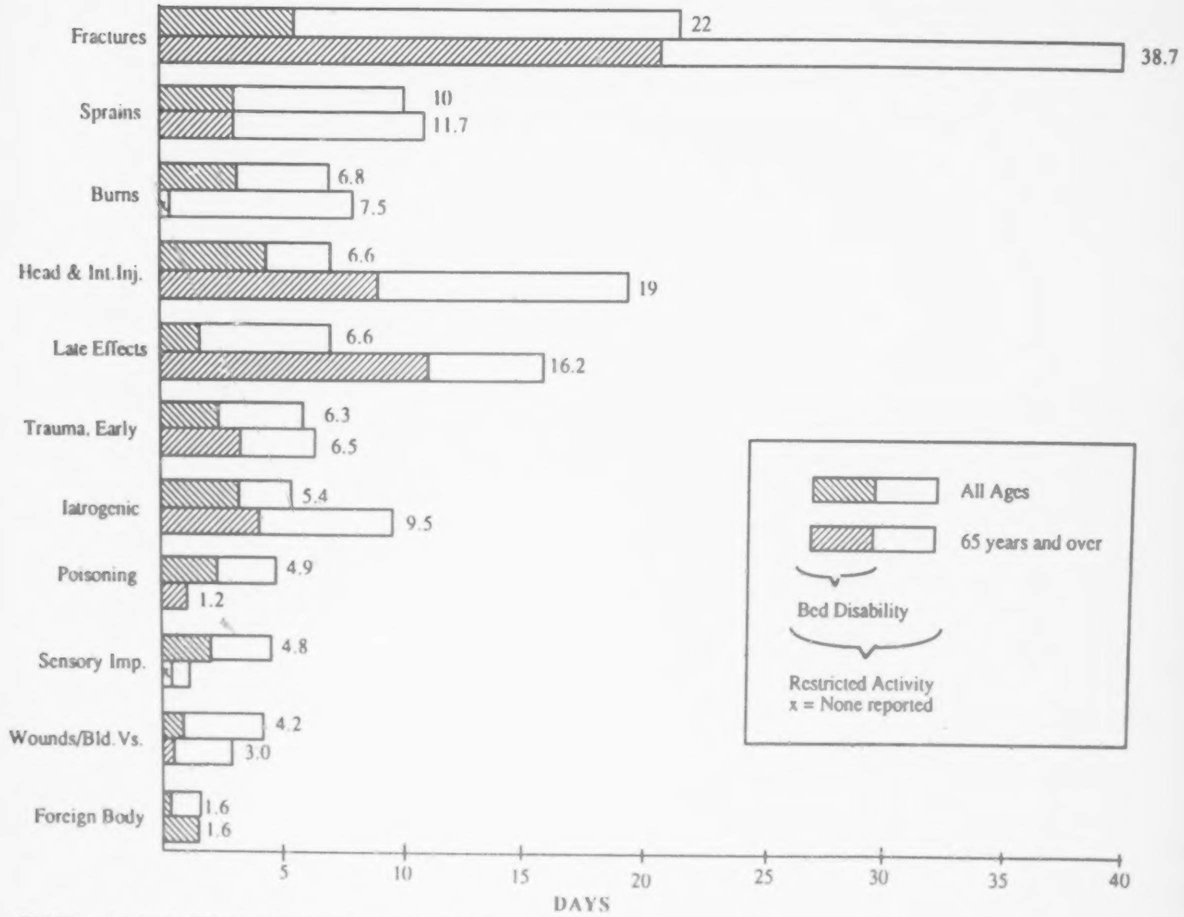
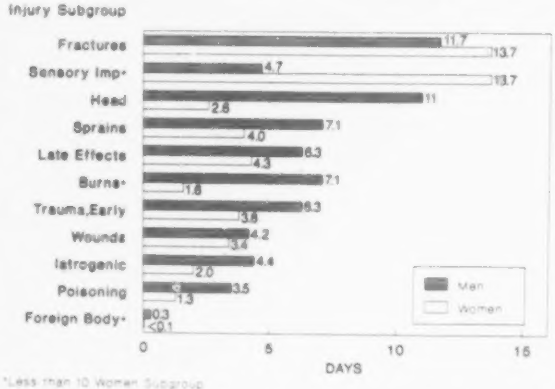


FIGURE 4—Restricted Activity and Bed Disability Days for All Ages and for Age 65+



*Less than 10 Women Subgroup

FIGURE 5—Mean Work Loss Days by Injury Subgroup and Gender

The burden of medical costs from injuries are also disproportionately greater for males, persons other than Black, and those with family incomes less than $5,000 (in 1980 dollars). Each of these three groups reports a greater proportion of health care charges attributable to acute injuries and their sequelae than do their counterparts. This maldistribution of costs may reflect differential exposure to potentially injurious situations in the workplace or community. Medical costs for the category of "White and other" might be slightly higher because of the inclusion of "others" in the category. However, the sample weights were computed for "Whites and others" as one category. The "others" category is small and therefore the effect on this category would be minimal.

The economic impact and the demographic pattern of expenditures are concordant with epidemiologic estimates of the health impact of injuries. The potential years of preretirement work life lost as a result of premature mortality from intentional and unintentional injuries was over 4 million years in 1982. This toll was roughly twice the years of potential work life lost from either neoplasms or heart disease.[2] In the National Household Interview Survey of 1980–81, a very high incidence of intentional and unintentional injuries were reported; about 75 million injuries were reported by 70 million persons.[13] The injury rates were higher for males than for females (40 million versus 30 million), with the greatest number (21 million) occurring in the 17 to 44 year old age group. Both the NHIS and the NMCUES agree on the high occurrence rate in the US population and on the gender and age distributions. These morbidity estimates also correspond to the patterns of national mortality from injuries.[14] Therefore, the mortality, morbidity, and cost data confirm the remarkable impact of injuries.

The economic burden of disease includes the loss of economic productivity to society and income resulting from work loss. These measures of lost productivity were high, reflecting the frequent occurrence of injuries among the younger, working-age population. Injuries accounted for approximately 10 percent of all restricted-activity and bed-disability days during the survey year and about 17 percent of all work-loss days. These estimates exceed the disability and work loss attributable to acute respiratory conditions, which are the most common acute conditions in the general population.[15] These findings agree generally with data from the larger National Health Interview Survey conducted in 1980–81.[13] These findings do not address additional mortality costs to society. Depending on the discount rate used, the mortality costs from injuries are reported to be six to eight times the morbidity costs.[12]

There are considerable social costs in terms of work loss, restricted activity, and bed disability; these costs are distributed across the entire age range. The duration of disability and the mean direct costs are greater for older persons, but the impact of total direct costs and work loss is nevertheless severe for younger adults. MacKenzie, *et al*, examined the functional recovery and medical costs of trauma in patients admitted to two trauma centers and discharged to extended care or rehabilitation facilities or home.[16] Of 262 individuals who were working before their accident, only 57 percent had returned to full-time work within one year.

Examination of total and per capita charges for diagnostic groupings provides further insights into the high costs of injuries. Hospital-based services, including physician fees, dominated the costs for all injury groups. This close linkage of injury care to hospital services indicates that costs for these conditions will parallel the general inflation of hospital operating costs. Because those with the least ability to pay—younger persons and those with low family incomes—have disproportionately higher costs for injuries, it is likely that increasing financial pressures may force many hospitals, and especially those with a large indigent or uninsured clientele, to reassess their ability to provide injury services.[17] A number of hospitals have already terminated emergency care services as a result of cost considerations. In a study of nonaccidental injuries, assaults and suicides, 2,451 patients were admitted to a trauma center.[18] The average charges for the victims of assaults were over $5,000 per patient while the suicide victim's charges were nearly $13,000. Six months after the study had been completed only about 70 percent of the hospital charges had been reimbursed. The authors noted that even if all of the hospital bills were covered the profit margin would only be 1 percent.

The charges associated with specific diagnoses identify conditions and characteristics of persons responsible for generating costs. Fractures accounted for the greatest total direct costs, and per capita costs were also high for this diagnosis. Moreover, fractures were responsible for greater mean disability- and work-loss days than were other diagnoses. Although fractures occurred more frequently in young adults (17 to 44 years), the cost and disability impact was greatest for those over age 65 years. However, fractures occurring in the older population are more likely to result from osteoporosis and hip fractures.[19,20] The diminished resiliency, more frequent requirement for surgical management, and the complicating comorbidity of chronic disease contribute to the greater burden for older persons. For similar reasons, head injuries and late effects of trauma also had a greater effect on older persons. These findings argue that injury research and preventive efforts should not be directed exclusively to the younger adult and employed population, who are generally considered the primary target for these efforts. Prevention and more efficient management of these injuries in the elderly could have considerable economic and health impact.

For conditions other than fractures, the aggregate charges and per capita charges were frequently discordant. For sprains, the per capita charges were relatively low but the high incidence in the population contributed to relatively high total charges. On the other hand, burns occurred less frequently but had higher per capita charges.

The diversity of the conditions included as injuries and the differences among them with respect to per capita costs and disability have implications for public policy. The most effective medical and cost reduction strategy would be prevention. Prevention would affect not only the less common and more expensive conditions but also the total costs of less expensive and commonly occurring conditions by decreasing the incidence of both. Prevention requires specification of the site and circumstances of injury. In NMCUES, there was no inquiry about the causes for the injuries or the sites of occurrence. However, in the contemporaneous National Health Interview Survey, vehicular accidents, industrial accidents, and the home were the most frequent sites of injury.[13] Examining type of accident or manner of injury as the cause of death for 1981, motor vehicle accidents were by far the leading cause of death,[21] followed by falls, fires, and drowning. Major cost containment strategies should be directed toward prevention at all potential sites as well as toward cost efficient urgent and emergency care. These data from NMCUES document the cost of post-emergency care for fractures, trauma, and burns and indicate the potential value of developing more cost-efficient care.

The remarkable impact of injuries raises important public policy issues. Compared with the considerable attention directed toward heart disease, cancer, and recently AIDS (acquired immunodeficiency syndrome), the health professionals and the public appear less concerned with injury research and prevention. The federal research investment in prevention and treatment of injuries was estimated to be $112 million in 1983 or about one-tenth of the research commitment to cancer and one-sixth of that to heart disease and stroke.[1] Moreover, only about one-fourth of federal research investment in injuries was directed toward prevention with the remainder devoted to injury mechanisms, acute care, and rehabilitation.[1] In the context of high medical care costs, this research investment is extremely small. The national costs for injury care exceed those for cancer and are nearly as great as for heart disease. The years of life lost and indirect costs from disability and work loss exceed those for either of these disease categories. The extremely small research commitment to injuries contrasts sharply with the commitment to cancer and heart research. From the perspective of research investment as a proportion of health care costs, injury had less than 1 percent (actually 0.7 percent) of the care cost invested in research. The comparable proportions for cancer and heart disease were 9 percent and 3 percent, respectively.[1]

Several philosophical, political, and financial reasons can be cited for the public oversight. Importantly, injuries may still be considered to result from random and unalterable events and thus be unpredictable and unpreventable. Even the relatively sparse epidemiologic data from investigation of accidents, injuries, and outcomes indicate that this is not true.[22] Further investigation should improve prediction and afford major decreases in injury rates and injury costs.[23] The lack of research monies probably relates to the absence of a focus for organizing a professional and public constituency. Without a focus of responsibility for defining and reporting injuries, funding research, and monitoring change, there is little political action to reassess priorities and generate new and useful knowledge. From the economic as well as the health perspective, the need warrants greater attention.

APPENDIX
Diagnostic Categories and ICD Codes for Accidents and Injuries

Diagnostic Category	ICD Code
Fractures	800–829
Sprains, strains, dislocations	830–839 840–848
Intracranial, internal injury	850–869, 950–957 special NCHS codes
Open wounds, injury to blood vessels	870–879, 880–887, 890–899, 900–904
Late effects of injuries, poisonings, toxic effects of external cause	905–909
Other early complication of trauma	910–919, 920–924 925–929, 958–959, 990–995
Effects of foreign body entering through orifice	930–939
Burns	940–949
Poisonings, toxic effects	960–979, 980–989
Complications of medical, surgical care	996–999
Partial impairment of senses or other special impairment from accident or injury	special NCHS codes for sensory impairment

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Evaluation of Active versus Passive AIDS Surveillance in Oregon

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Abstract: We searched for unreported AIDS cases in Oregon through death certificate and medical record review, and enhanced infection control practitioner and physician surveillance. Fifty-six AIDS cases diagnosed between February 1, 1986 and January 31, 1987 were reported passively. Twenty-nine additional cases diagnosed during this time were retrospectively identified by active methods. Ninety percent of those 29 cases were diagnosed by physicians and cared for in hospitals that had previously reported cases. Completeness of reporting under the passive system was 64 percent. (*Am J Public Health* 1990; 80:463–464.)

Introduction

Since the acquired immunodeficiency syndrome (AIDS) was first described in 1981, surveillance for AIDS cases has served to define the modes of transmission and to characterize the populations most affected. To date, there have been little data to assess the need for these active surveillance systems, that seek AIDS case reports through a variety of methods rather than passively depending on physician reporting. This study assesses the need for active AIDS surveillance systems. The HIV (human immunodeficiency virus) program of the Oregon Health Division initiated prospective active surveillance for AIDS in February 1987. We applied the methods used for active surveillance to retrospectively determine the number and characteristics of AIDS cases that had gone unreported under the previous passive system. We report here the results of that evaluation.

Methods

We used four active surveillance methods to identify unreported cases diagnosed between February 1, 1986 to January 31, 1987, the year before active surveillance began.

- All physicians in Oregon were sent a letter reviewing AIDS reporting requirements and soliciting case reports. Physicians who had previously reported at least one AIDS case were provided with a list of all these cases for review and asked to report additional unreported cases under their care.
- Infection control practitioners at hospitals where previous AIDS cases had been diagnosed were asked to compile a list of previous AIDS admissions and maintain a log of subsequent admissions for comparison with our central AIDS registry. Those from hospitals without reported cases were asked every six months to confirm that no AIDS-related admissions had occurred.
- At all hospitals statewide where at least one case had been diagnosed, medical records of all potential AIDS admissions were reviewed. Admissions were pulled if they contained any of five International Classification of Diseases, Volume 9 (ICD-9) codes (042, 043, 279.19, 136.3, 173.9) previously shown to be reliable in identifying persons with an AIDS diagnosis.[1] Cases were matched with reported cases in the central registry and physicians of possible unreported cases were queried.
- Death certificates designating AIDS or listing a cause of death potentially AIDS-related (including all deaths related to known opportunistic conditions and deaths from infectious or immunosuppressive causes in persons under age 60) were reviewed. Physicians of potential unreported cases were queried.

A case diagnosed during the year before active surveillance began was considered to be discovered by **active surveillance** if it was reported by any reporting source after the start of active surveillance and more than 6 months after diagnosis. All other cases were classified as reported by passive surveillance. All cases met the 1985 CDC surveillance case definition.[2]

Results

Completeness of Reporting

Fifty-six AIDS cases diagnosed between February 1, 1986 and January 31, 1987 were reported under the passive system. Using active surveillance methods an additional 29 cases diagnosed during this time were found retrospectively. The completeness of reporting before active surveillance began was 64 percent (95% confidence interval = 54, 74). For cases reported under the passive system, the median time between diagnosis and report, or lag time, was one month. Lag time ranged from 7 to 19 months for cases found by active surveillance only (median = eight months).

Characteristics of Cases and Reporting Sources

Cases reported by passive surveillance did not significantly differ from those identified by active surveillance by age, race, county of residence, risk factors, disease at diagnosis or mortality status at time of report, although our power to detect significant differences was limited by the relatively small sample size (Table 1).

Characteristics of physicians who reported cases by passive surveillance did not differ from physicians whose cases were identified under active surveillance by specialty,

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TABLE 1—Characteristics of Cases Reported by Passive or Active Surveillance

Characteristics	Passive (n = 56)		Active (n = 29)	
Age (mean)	35.6yrs		35.2yrs	
Race	No	%	No	%
White	53	94.6	28	96.6
Non-White	3	5.4	1	3.4
County of residence				
Multnomah (urban Portland)	38	67.9	23	79.3
Other	18	32.1	6	20.7
Mortality status				
Alive at report	38	67.9	19	65.5
Dead at report	18	32.1	10	40.0
Disease at diagnosis				
PCP	33	59.1	20	68.5
Other disease w/o PCP	14	25.2	3	10.3
KS alone	9	16.1	6	20.4
Mode of transmission				
Homosexual/bisexual men	45	80.4	22	75.9
Homosexual/bisexual IVDA	5	8.9	5	17.2
Hemophiliac (adult)	3	5.4	1	4.0
Other	3	5.4	1	4.0

PCP = Pneumonocystis carinii pneumonia; KS = Kaposi's sarcoma; IVDA = intravenous Drug Abuse

city of practice, or previous reporting history. Ninety percent (26/29) of cases identified by active surveillance were cared for by physicians who had reported at least one case before the start of active surveillance.

Characteristics of hospitals from which cases were reported passively did not differ from hospitals where cases were identified under active surveillance by previous reporting history, bed size, or location. Ninety percent (26/29) of cases identified by active surveillance were diagnosed in hospitals that had cared for at least one case before the start of active surveillance.

Efficacy of Active Surveillance Components

We compared the relative effectiveness of each component of active surveillance in identifying previously unreported cases. The initial reporting source for cases identified actively was physicians for 14 percent, infection control practitioners 31 percent, medical record review 31 percent, and death certificate review 24 percent. We estimated the completeness of the surveillance system using the capture-recapture methodology (Chandra Sekar and Deming method).[3] Our surveillance system was divided into two independent reporting components: unsolicited (physician or infection control practitioner reporting) and solicited (medical record review or death certificate review). The percentage of cases found by unsolicited means was 80 percent. The percentage of solicited reports was 95 percent. Seventy-seven percent of cases were found by both methods. The estimated completeness of reporting was 99 percent of the estimated case population.

Discussion

Relative to active AIDS surveillance, the completeness of passive AIDS surveillance in Oregon in 1986–87 was only 64 percent, similar to preliminary findings from South Carolina[4] and less than an earlier New York study.[5] Passive AIDS surveillance, even in relatively low incidence areas, is incomplete. Projections based on passive surveillance would underestimate future incidence proportionately and could result in poor public health planning and resource distribution decisions.

TABLE 2—Characteristics of Physicians Reporting by Passive or Active Surveillance

Physician Characteristics	Passive		Active	
Specialty	No.	%	No.	%
Internal Medicine	29	51.8	10	36.0
Infectious Disease	10	17.9	13	44.8
Allergy and Immunology	9	16.1	1	3.4
Other	6	10.7	5	17.2
Unknown	2	3.6	0	0
City of practice				
Portland	39	69.6	24	82.8
Not Portland	17	30.4	5	17.2
Previous AIDS Report				
Yes	42	75.0	26	89.7
No	13	23.2	3	10.3
Unknown	1	1.8	0	0

Because almost all cases during this time in Oregon were occurring in urban gay and bisexual men, we recognized at the outset that we would not be able to assess differences in reporting based on case risk factor.

Infection control practitioners, medical record review, and death certificate review were all useful in identifying cases unreported more than 90 days after diagnosis. Each of the multiple reporting components independently identified cases that would otherwise have gone unreported.

Physicians and hospitals with previous reporting history appear to be the reservoir for the largest number of unreported cases and are the best targets for enhanced surveillance activities. As the epidemic continues, the ability to maintain effective active surveillance will include using resources efficiently, combating "reporting fatigue," and assuring timely reporting as clinical practice shifts diagnosis to outpatient settings. As the AIDS epidemic progresses and physicians and hospitals gain more experience with this disease, active surveillance may become more rather than less important.

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The Prevalence of High-Risk Sexual Behavior in Male Intravenous Drug Users with Steady Female Partners

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Abstract: A sample of 149 (70 White, 79 Black) male intravenous drug users with steady female sexual partners was interviewed in 1987, in treatment and street settings, about sexual practices during the previous five years. Eighty-three percent had multiple partners, 15 percent reported male sexual contact, 38 percent reported heterosexual anal intercourse, and 73 percent never used condoms. Thirty-seven White and 30 Black subjects reported bisexuality or heterosexual anal intercourse or both. (*Am J Public Health* 1990; 80:465–466.)

Introduction

Research on AIDS (acquired immunodeficiency syndrome) prevention among intravenous (IV) drug users has focused primarily on needle sharing,[1,2] with few studies to date on high-risk sexual behavior.[1,3] Since sexual behavior may prove more difficult to change than needle behavior,[2,4,*] more information is needed to design effective AIDS education programs. This paper presents the prevalence of high-risk sexual behavior in a sample of male IV drug users in San Francisco.

Methods

Data on sexual behavior were collected in early 1987 from 149 male IV drug users with stable female sexual partners. The men were a sub-set of a larger study of risk behaviors for human immunodeficiency virus (HIV) transmission among 632 male and female IV drug users. Of the 411 male subjects, 34 had exclusively male sexual contacts during the previous five years and 377 had sexual contact with females during that period. Of these 377, 134 were White, 171 were Black, and 82 were Hispanic or other ethnic backgrounds. Participants were asked: "Do you have a steady sexual partner (or partners)?" If they answered yes, they were asked whether the partner(s) was male, female, or both. Of the 377 males, 201 reported a steady partner or partners; 184 had a steady female partner, 14 had a steady male partner, and three had both a male and a female partner. One White and one Black male reported both a female and a male as steady partners. Information on length of relationship was not obtained. Men with no steady sexual partners, those whose partners were exclusively male, and those of Hispanic or other ethnic backgrounds were not included in this study.

Subjects were recruited from a 21-day methadone clinic and a symptomatic (nonmethadone) clinic, and from three street settings. In-treatment subjects had a documented history of opiate addiction (the methadone clinic), or were identified by staff as having injected drugs during the three weeks prior to admission (the symptomatic clinic). Of those contacted in the two clinics, 86 percent agreed to participate. Out-of-treatment subjects, recruited via word of mouth through a targeted sampling technique,[5] reported needle use within three weeks of interview and showed signs of recent venipuncture. The refusal rate for street subjects is unknown because of the chain-referral recruitment method.

Structured 45 minute interviews were conducted by trained interviewers who questioned subjects about demographic characteristics, medical history, drug use patterns, and AIDS awareness, as well as sexual behavior. Subjects were asked their total number of female sexual partners during the previous five years; if they had sexual contact with males; if they ever engaged in heterosexual anal intercourse, and the percent of time of condom use.

Results

Of the 149 subjects, 70 were White and 79 were Black; less than one-fifth (18 percent) of the subjects were currently married. Age ranged from 20 to 66 years; the median age of Whites was 33.5 years, of Blacks 36 years. Eighty-eight (59 percent) of the men were interviewed out-of-treatment, 27 (39 percent) of the White subjects and 61 (77 percent) of the Black subjects.

High Risk Sexual Activities

Over four-fifths (83 percent) of the men reported more than one female sexual partner during the previous five years, 60 percent had five or more, and 39 percent had 10 or more. Whites had a median of 7.5 partners and Blacks a median of five partners.

Twenty-two (15 percent) of the 149 men reported sexual contact with a male (13 percent of the Whites and 17 percent of the Blacks). White subjects who reported same-sex activity had a median of 20 female partners and three male partners. Black subjects who reported bisexual activity had a median of 10 female and three male partners.

Heterosexual anal intercourse was reported by 56 (38 percent) of the men (46 percent of the Whites and 30 percent of the Blacks). Almost three-fourths (73 percent) of the men said they never used condoms.

Table 1 shows the distribution of subjects engaged in high-risk sex by treatment status and by ethnicity. Twenty-seven of 61 in-treatment subjects and 29 of 88 out-of-treatment subjects reported either bisexuality or heterosexual anal intercourse. In all, 33 White and 23 Black subjects reported either male sexual contact or heterosexual anal intercourse; seven Black and four White subjects engaged in both activities (see Table 1).

The majority of men engaged in high-risk sex reported non-use of condoms (Table 1). Fourteen of 22 men who had male sexual contact, and 41 of 56 men who engaged in

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TABLE 1—High-Risk Sexual Behaviors in 149 Intravenous Drug Users According to Treatment Status and Ethnicity

	(1)	(2)	(3)		(4)
	Bisexual	Heterosexual Anal	(1)+(2)	Either Heterosexual Anal or Bisexual	Never Use Condoms
In-Treatment (N)					
Black (18)	1	9	1	8	16
White (43)	3	20	2	19	33*
Total (61)	4	29	3	27	49*
Out-of-Treatment (N)					
Black (61)	12	15	6	15	38
White (27)	6	12	2	14	21
Total (88)	18	27	8	29	59
Total (N)					
Black (79)	13	24	7	23	54
White (70)	9	32	4	33	54
Total (149)	22	56	11	56	108

*One subject with a missing answer.

heterosexual anal intercourse said they never used condoms (data not shown).

Discussion

Overwhelmingly research on IV drug users is conducted in treatment programs.[5,6] In this study, the majority of subjects were recruited out-of-treatment. Because this was not a random sample, the results may not be generalizable. Other limitations are a lack of information on duration of partner relationships and the problem of bias in recalling and accurately portraying past sexual activity.

Due to a lack of comparable studies, it is unknown whether the range and level of sexual activity reported by these subjects is similar to other IV drug users. The sexual risk taking in this sample may be connected with recent increases in cocaine injection and smoking.[7,8] Because cocaine has been linked with heightened sexual activity in some studies,[9,10] the impact of it and other drugs on high-risk sexual behavior merits study.

The numerous sexual partners and the high-risk practices reported by a majority of these IV drug users indicate the potential for widespread secondary transmission of HIV. The difficulty in reducing sexual risk behavior is shown by the high proportion of subjects in this study who reported non-use of condoms, despite out-reach risk reduction programs, initiated in 1986, which included widespread distribution of free condoms to IV drug users in San Francisco.[4,**]

Further research is required on ways to overcome barriers to condom use and to reach heterosexually identified male IV drug users engaged in sexual activity with males. Continued aggressive culturally sensitive out-reach sexual risk reduction strategies are needed, since many IV drug users may not be enrolled in treatment programs.[11–13]

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Are Patients Talking to Their Physicians about AIDS?

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Abstract: We conducted a nationwide telephone survey of a random sample of United States adults in summer 1988 (n = 2000, response rate = 75 percent) to find out if physicians were providing education and counseling to the public about AIDS and AIDS prevention. Within the previous five years, 94 percent had seen a physician but only 15 percent had discussed AIDS even though most said they would not object to discussing the topic. AIDS-related conversations are not commonplace in physician's offices and in most cases (72 percent) patients are the initiators of such conversations. (*Am J Public Health* 1990; 80:467–468.)

Introduction

Physicians have been encouraged to discuss AIDS (acquired immunodeficiency syndrome) risk reduction with all of their patients and to take sexual histories to help all patients identify and modify risky behavior;[1,2] physicians are well placed to provide this service, as approximately 76 percent of Americans see a physician annually.[3] Moreover, 60 percent of respondents to a National Center for Health Statistics survey[4] said that they would ask a physician if they wanted specific information about AIDS.

Methods

Telephone interviews were completed with a representative sample of 2000 civilian, non-institutionalized United States adults (75 percent response rate) between July 5 and August 19, 1988. In selecting the sample, the country was divided into two strata by AIDS prevalence and subjects were drawn using a random-digit-dial procedure. There were 1000 respondents from each stratum.

The data reported here are based on weighted estimates for the two strata combined. Given the stratification design effect (the ratio of the variance of design used to that of a simple random sample of the same size), the effective sample size for population estimates was 1320.[5] The confidence intervals in Table 1 reflect this design effect.

Subjects were selected randomly within households. To partially correct for known biases in the selection of respondents resulting from random digit dialing,[6] responses were weighted by the number of persons in the household and the inverse of the number of telephone lines in the household. To correct for small differences between the demographic characteristics of the sample and census estimates, data were also weighted by race, sex, and age.

The interview contained 70 items concerning public perceptions about AIDS and health care. Items about physician counseling are presented in the Appendix. Ninety-four percent of respondents recalled seeing a physician at least once in the past five years; 76 percent of respondents had visited a physician in the past year. Data reported here are from respondents who had seen a physician in the past five years, referred to later in this paper as "patients." Further details of the methods have been reported in earlier papers.[7-9]

TABLE 1—Percent of Patient Respondents Who Had Discussed AIDS with their Physician by Demographic Characteristics (with 95% confidence intervals)

	N*	%	95% Confidence Interval
Total	1876	15.4	12.9, 17.9
Sex			
Male	865	14.1	11.2, 17.0
Female	1011	16.4	13.6, 19.2
Race			
White (non-Hispanic)	1496	14.5	12.3, 16.7
Black	180	22.2	14.7, 29.7
Age (years)			
18–29	511	17.2	13.2, 21.2
30–49	789	17.5	14.2, 20.8
50+	576	10.8	7.7, 13.9
Income (household per annum)			
< $20,000	478	13.0	9.3, 16.7
$20,000–34,999	535	13.8	10.2, 17.4
$35,000–49,999	320	16.3	11.3, 21.3
$50,000+	220	20.9	14.3, 27.5
AIDS Prevalence in SMSA			
High	943	19.4	16.9, 21.9
Low	933	11.3	9.3, 13.3
What are your chances of getting the AIDS virus?			
High/Medium	96	27.1	16.2, 38.0
Low	660	17.4	13.8, 21.0
None	1090	12.8	10.4, 15.2
Knows HIV-infected person	372	26.9	21.4, 32.4
Does not know HIV-infected person	1474	12.4	10.3, 14.5
Knows risk group member	761	23.4	19.7, 27.1
Does not know risk group member	1050	10.0	7.8, 12.2

*Subtotals which do not add to 1876 exclude "don't know" and "no answer" respondents for those items.

Results

Only 15 percent of patients recalled ever talking to their physicians about AIDS. The demographic groups more likely to have discussed AIDS were young people, Blacks, people with a high household income, or people who lived in an SMSA with high AIDS prevalence (Table 1). In addition, those for whom AIDS was personally most salient, for example those who knew someone with AIDS or HIV (human immunodeficiency virus) infection, were more likely to have talked to their physician about the disease (Table 1). Of the 240 patients who had talked to their physician about AIDS, physicians initiated the conversations in only 28 percent of the cases.

Patients had discussed a variety of AIDS-related issues with their physicians; sexual behavior and safe sex were discussed most frequently. Risk of transmission from blood transfusions

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and other health care-related risks were also discussed frequently, as were concerns about casual transmission.

Among patients who had not yet spoken with their physician about AIDS, 56 percent did not express a preference for or against talking to their doctor about AIDS, only 17 percent said they would not want to discuss the topic with their doctor, and 27 percent wanted to do so. Younger respondents (32 percent) and Blacks (37 percent) were more likely to want to talk to their physician about AIDS.

Since many AIDS-related concerns have to do with sex, patient-physician interaction regarding sexuality may influence the frequency of discussions regarding AIDS. Most patients (78 percent) did not recall having been asked questions about their sexual behavior at their physician's offices. Only 7 percent of respondents reported they would be unwilling to answer a physician's questions about their sexual behavior; 59 percent said they would be glad for the opportunity to talk to their physician about sex. Most patients (67 percent) said they would be very comfortable talking with their physician about AIDS.

The AIDS-related topics which patients said they would like to discuss are listed in Table 2. Transmission of AIDS, especially by casual contact and in medical settings, heads the list. Although 59 percent of patients said that they would go to their physician or HMO (health maintenance organization) for more specific information on AIDS, respondents reported getting most of their information from television (mentioned by 67 percent), newspapers (56 percent), and magazines (36 percent).

Discussion

Relatively few patients had discussed AIDS with their physician. This is consistent with previous reports based on physician surveys[10] and more limited patient surveys.[11] Previous studies[12] have also documented the infrequency of sexual history taking. It should be noted, however, that the denominator for the history taking and AIDS counseling items includes all patients who have had any physician contact over the past five years. Many of these contacts may not have been appropriate situations for these preventive practices. Although fallibility of patient recall may also lead patients to underestimate counseling by physicians, if patients do not remember having discussed AIDS, it is unlikely that the counseling was effective. Although the methodology of this survey may yield conservative estimates of physicians' actual counseling efforts, the evidence remains that AIDS-related discussions are not commonplace in physicians' offices.

TABLE 2—AIDS Topics People Wanted to Discuss with Their Physician according to the 419 Patient-respondents (27%) Who Said They Wanted to Talk to Their Physician about AIDS but Had Not Yet Done so

Topics for Discussion	%
AIDS transmission—general information	23.9
Risk from casual contact	21.0
Contagion in medical settings	11.7
Sexual behavior, safe sex	9.5
Prevention/avoid contracting AIDS	7.9
Treatment for AIDS	7.6
Infection control	6.7
Blood transfusions	5.7
Symptoms of disease	5.7
If doctor knew of anything not widely reported	5.3
Antibody testing	3.8
If physician was treating people with AIDS	3.6
Scientific questions	3.3
Drug use	3.1
Distrust of mass media	2.9

APPENDIX

Survey Items Concerning Patients' Experience and Opinions about AIDS Education by Physicians

- Have you ever talked to your medical doctor about AIDS?
- Who started the conversation . . . you or the doctor?
- What was it about AIDS that you and your doctor discussed? *(open-ended)*
- Would you want to talk to your doctor about AIDS, would you prefer not to discuss this with your doctor, or do you not care?
- What is it about AIDS that you would like to discuss with your doctor? *(open-ended)*
- If you were to talk with your doctor about AIDS, should they start the conversation, or would you rather start it yourself?
- Have you ever been asked questions at the doctor's office about your sexual behavior?
- It has been suggested that medical doctors should ask their patients about their sexual behavior as part of a routine health history taking. If your doctor asked you about *your* sexual behavior, would you be glad for the opportunity to share this information with your doctor, be somewhat uncomfortable but answer the questions anyway, or would you be unwilling to talk about these matters with your doctor?
- How comfortable would you feel . . . *(insert a through c)* . . . very comfortable, mostly comfortable, somewhat uncomfortable, or very uncomfortable?
  a. if you had a conversation with your doctor about AIDS
  b. if your doctor asked you questions about your sexual behavior
  c. if your doctor asked you questions about your use of intravenous drugs
- According to medical authorities, the following groups of adults are at highest risk for AIDS: people who received blood transfusions between 1977 and 1985, intravenous drug users, men who have sex with other men, and anyone who has had many sex partners. Do you know anyone that might be a member of these groups?
- Have you ever personally known anyone with AIDS or the AIDS virus?

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Screening for Chlamydial Cervicitis in a Sexually Active University Population

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Abstract: Enzyme-linked immunoabsorbent assays to detect chlamydial cervicitis were performed on samples from 1,320 sexually active university women. Seventy-five (prevalence 5.7 percent) had positive tests. Demographic, history, symptom, and physical examination variables were insufficient to predict infection accurately. We conclude that screening during routine visits with this population is cost-effective. (*Am J Public Health* 1990; 80:469–471.)

Introduction

Chlamydia trachomatis is the most prevalent sexually transmitted disease in the United States.[1] Until recently, diagnosis of chlamydial infection was handicapped by the lack of an inexpensive diagnostic test. *Chlamydia* is easily treated with antibiotics, but left untreated can persist for an indefinite time, and result in severe sequelae. While symptomatic women have a higher prevalence rate, most women with cervical infections have few or no symptoms.[2]

Methods

Routine screening for chlamydial infection of all sexually active students receiving physical examinations at the University of California-Los Angeles, Student Health Service Women's Clinic was implemented in June 1986. Some women were seen for initial or yearly examinations (routine visit group) and some women were seen for symptoms, partner symptoms, or previous problem follow-up (symptom/problem group).

Specimens were tested using Chlamydiazyme™ kits run according to the manufacturer's directions. Women with positive tests were typically tested with doxycycline 100 milligrams twice daily for seven days. Sexual abstinence was encouraged during treatment; partners were referred for treatment. Test of cure using culture was requested approximately three weeks after treatment.

A total of 1,320 women were screened between June 24 and October 24, 1986. Data were abstracted from the medical records of all but 19 of these women. Data collected included: age, marital status, race/ethnicity, reason for the visit, gynecologic history, present contraceptive use and level of current sexual activity, current symptoms, partner symptoms, physical findings, test results, and clinician diagnosis.

Chlamydiazyme test result was cross tabulated with these variables; prevalence odds ratios (POR) with 95% confidence intervals (CI) were calculated. All variables found to have POR with a lower limit on the 95% CI greater than one were subsequently used in a logistic regression (BMDPLR).[3]

Results

Of the 1,320 women tested, 75 (5.7 percent) had positive Chlamydiazyme tests. Thirty-six (5.0 percent) of the 724 routine visit group women had positive Chlamydiazyme tests while 39 (6.8 percent) of the 577 symptom/problem group women had positive tests.

Seventy-four of the 75 women with positive tests at their first study period visit returned for treatment; 65 of the 74 treated women returned for a culture test of cure. All 65 were successfully treated.

Table 1 lists those variables which were related to Chlamydiazyme result for the total sample. In the logistic regression, only five of the 13 variables were independently related to a positive test.

Recognizing that a logistic regression equation is likely to be too complicated for most clinicians to apply, a risk-factor index similar to those used by others[4,5] was constructed by summing presence of positive findings on each of the five independently related variables. Thus, women with none of the findings were assigned a score of 0, and women with all five of the findings were assigned a score of five. Table 2 lists sensitivity, specificity, and predictive values using different scores on the index as cutpoints.

Parallel analyses to those conducted for the total sample were conducted separately for women who were seen for routine visits, and those who were seen for problem/symptom visits. Results were not substantially different, but are available upon request

Discussion

Prevalence rates of chlamydial infection among female college students have ranged between less than 5 percent to nearly 10 percent.[6–14] The 5.7 percent prevalence of cervical infection we observed is at the low end of this range. However, our prevalence among women age 22 years and younger in our study was 8.7 percent; the relatively large number of older students in our study may account for the lower rate.

The literature is divided over whether medical history and examination are adequate to identify infection.[12,14] In our study, if only women who had one or more of the risk factors found to be independently related to a positive test had been screened, 694 women would not have been screened at a significant savings in resources, and 81 percent of the women

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TABLE 1—Rates of Positive Chlamydiazyme Tests Related to Selected Demographic, History, Symptom, and Examination Variables

	All Subjects		
Variables	N	Percent Positive	Prevalence Odds Ratio (95% CI)
Marital status			
Not married	1096	6.1	2.7
Married	129	2.3	(0.8,8.8)
Age (years)			
22 or younger	462	10.0	3.1
23 or older	839	3.5	(1.9,5.0)**
Sexually active past 3 months			
No or not noted	494	3.6	2.0
Yes	807	7.1	(1.2,3.5)
New partner past 3 months			
No or not noted	1163	5.0	2.7
Yes	138	12.3	(1.5,4.9)**
Multiple partners past 3 months			
No or not noted	1229	5.2	3.3
Yes	72	15.3	(1.7,6.6)
Taking oral contraceptive			
No	831	4.7	1.7
Yes	470	7.7	(1.1,2.7)
Using condoms			
No	1164	6.2	2.9
Yes	137	2.2	(0.9,9.4)
Partner present symptoms			
No or not noted	1240	5.1	4.6
Yes	61	19.7	(2.3,9.1)
Cervix normal on examination			
No	77	15.6	3.4
Yes	1218	5.2	(1.7,6.8)
Cervix bleeds easily			
No	1263	5.3	6.0
Yes	32	25.0	(2.6,13.7)**
Cervix red			
No	1267	5.5	3.9
Yes	27	18.5	(1.4,10.6)
Wet mount—yeast			
No	1090	6.3	2.3
Yes	209	2.9	(1.0,5.4)
Wet mount—bacterial vaginosis			
No	1217	4.8	5.2
Yes	82	20.7	(2.9, 9.4)**
Wet mount—white blood cells			
No	1254	5.2	5.2
Yes	45	22.2	(2.5,10.7)
Diagnosis—normal examination			
No	621	7.7	2.0
Yes	678	4.0	(1.2,3.2)
Diagnosis—possible chlamydia			
No	1265	5.0	10.4
Yes	34	35.3	(5.0,21.5)**

**Variables independently related to Chlamydiazyme result in logistic regression

with a positive test would have been tested. However, 14 (19 percent) of the 75 women with positive tests would have been missed. In our sample, 27 (36 percent) of the 75 women with positive tests had entirely normal physical examinations. Thus it appears while signs, symptoms, history, and demographic variables would be helpful in determining whether to test for infection when laboratory resources are limited, these variables are insufficient to predict infection *accurately*.

Seven guidelines for determining whether a screening program is likely to be effective have been proposed.[15] If a randomized clinical trial has shown the effectiveness of a screening program, the assessment is relatively easy. We are unaware of any randomized trials of screening for chlamydial infection. However, we believe a screening program in our female population satisfactorily satisfies the other six guidelines.

For example, several tests are available which could be used to screen for *chlamydia*. We chose to use Chlamydiazyme, which has compared favorably with culture in our population.[10] The importance of the screening test and the prevalence of the condition is especially important when considering possible negative effects of a screening program. Using the specificity (99.5 percent) and prevalence (4.9 percent) found in our previous studies,[10] less than one out of 10 women who are told they are infected will be erroneously told so.

The costs of alternative strategies for screening during routine gynecologic visits were recently analyzed.[16] A strategy using no screening was compared with a strategy involving the routine use of culture, and a strategy using a less costly but less sensitive and specific alternative test. The analysis indicated that the alternative test would need to cost about $10 per test at a prevalence of 5 percent for the screening strategy to be cost effective. Because our screening test is both more sensitive and specific than the estimates used in that analysis, and our cost per test is less than $8, it appears that screening among our population can be supported on economic as well as clinical grounds.

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TABLE 2—Prediction of Positive Chlamydiazyme Result Using Different Levels of Risk Factor Index—All Subjects

Risk Factor Index Score	N	% Positive	% Sensitivity*	% Specificity*	% Positive Predictive Value*	% Negative Predictive Value*
0	694	2.0				
1 or More	599	10.2	81.3	55.8	10.2	98.0
0 or 1	1172	4.3				
2 or More	121	20.7	33.3	92.1	20.7	95.7
2 or Less	1270	5.0				
3 or More	23	47.8	14.7	99.0	47.8	95.0

*Compared to Chlamydiazyme result—predictive values based on observed prevalence

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Prevalence of Markers for Hepatitis B and Hepatitis D in a Municipal House of Correction

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Abstract: Following an outbreak of hepatitis B (HBV) in a municipal house of correction, HBV markers were detected in 173/406 (43 percent) inmates and 10/129 (8 percent) staff. Of the 173 HBV-infected inmates, 14 (8 percent) had hepatitis D (HDV) markers compared to 0/10 staff members. Intravenous drugs use (IVDU) was most strongly associated with HBV marker presence. Increasing duration of imprisonment, history of hepatitis B and especially IVDU were associated with the prevalence of HDV markers. (*Am J Public Health* 1990; 80:471–473.)

Introduction

Hepatitis D (HDV) is caused by an incomplete RNA virus dependent on hepatitis B virus (HBV) for survival.[1,2] Estimates of its prevalence range from 0–10 percent in chronic HBV carriers to 30–50 percent in persons with fulminant hepatitis B.[3] Most data suggest that in the United States intravenous drug users and hemophiliacs are at increased risk for HDV.[4] Although higher prevalence of HBV markers in prisoners or inmates of correctional facilities when compared to the general population has been described in several studies, similar data on HDV in incarcerated populations are not available.[5–10]

In January 1985, the Boston Department of Health and Hospitals was notified that five inmates in the Suffolk County House of Correction had an acute onset of hepatitis B. One inmate was coinfected with HDV and died of fulminant hepatitis. All inmates and employees were offered HBV screening and subsequent vaccination if seronegative as part of an outbreak control plan. While providing epidemiological assistance, we had the opportunity to evaluate the HDV status in those with HBV markers.

Methods

The Suffolk County (Massachusetts) House of Correction at Deer Island, operated by the City of Boston Penal Institution Department, accommodates 460–480 inmates, who serve an average sentence of 11 months. There are approximately 210 staff members employed at the facility including two registered nurses and one part time physician who staff an on-site infirmary.

In February 1985, a fact sheet on HBV and HDV was distributed to all inmates and staff, and participation in an HBV screening and vaccination program was solicited. Participants were questioned by one of the investigators regarding age, duration of involvement with the prison system, history of prior hepatitis, history of intravenous drug use (IVDU), location within the prison facility, and primary city of residence.

Serum was obtained to determine presence of HBV core antibody (anti-HBc; CORAB, Abbott Laboratories, N. Chicago, Illinois). Sera with borderline anti-HBc results were also tested for hepatitis B surface antigen (HBsAg) and antibody (anti-HBs) (AUSRIA and AUSAB, respectively, Abbott Laboratories). All anti-HBc positive specimens were also analyzed for HDV antibody (anti-delta; Abbott Laboratories). An interpretation of test results was given to each participant and HBV vaccination offered to those without HBV markers. This protocol was approved by the Boston Department of Health and Hospitals Institutional Review Board.

The magnitude of relative effect was assessed for several risk factors of HBV positivity. Univariate and multivariate

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TABLE 1—Characteristics of Participants in a Hepatitis Screening Program in Prison, Boston, 1985

Characteristics	Inmates (N = 406) (%)	Employees (N = 129) (%)
Mean age (years)	27.8	33.9
Mean duration of prison involvement (years)	2.2	5.7
History of hepatitis, any type	45 (11)	5 (4)
History of hepatitis B	17 (4)	0 (0)
History of homosexual activity	12 (3)	0 (0)
Intravenous drug use	136 (33)	2 (2)
Anti-HBc positive	173 (43)	10 (8)

analyses were conducted using SAS procedures to calculate odds ratios, test-based confidence intervals,[11] and logistic regression coefficients. Crude odds ratios (OR) and their exact 95% confidence intervals (CI) for risk factors in the 14 cases of HDV were calculated using the method described by Gart.[12]

Results

Of 463 inmates offered screening, 409 (88 percent) agreed to participate. Of these 409, three had equivocal anti-HBc results and are not included in further analysis. Characteristics of participants are shown in Table 1. There was no discernible pattern in findings related to city of residence.

Of the 142 employees invited to participate, 129 (91 percent) agreed. Their characteristics are shown in Table 1.

By univariate analysis, risk factors for HBV among prisoners included age greater than 30 years, prior history of HBV, and IVDU. In the multivariate model prior history of hepatitis was also included. Tests for interaction of age and IVDU and for other potential interactions were not significant. By multivariate analysis, IVDU was most strongly associated with HBV marker presence (OR = 6.0, 95% CI = 3.6, 10.2). Increasing duration of incarceration, independent of age or IVDU, was also related to the presence of HBV markers (Table 2).

Of the 173 specimens from inmates with anti-HBc, 14 (8 percent of all inmates) contained anti-delta. Relative odds of IVDU, imprisonment for one to four years and for five years or more, history of hepatitis B, and age over 30 were elevated. All prisoners positive for HDV reported that they had had some type of hepatitis in the past, and 12 of the 14 reported both a history of hepatitis and IVDU. The small numbers and correspondingly wide intervals, and the inability to control joint confounding in this data set, prevents further interpretation of the data.

Discussion

Several studies have documented a high prevalence of HBV markers in prison populations.[5–10] Reported figures are comparable to the proportion reported here of 43 percent. As in other studies, IVDU was found to be strongly associated with evidence of HBV infection. One-third of our inmate participants gave a history of intravenous drug use. This figure may not be reflective of correctional institutions as a whole, and thus the prevalence of HBV markers may differ in other facilities.

Less information is available on the prevalence of HDV infection in inmates of correctional facilities. A previous study revealed that four of 19 prisoners (21.1 percent) known

TABLE 2—Prevalence of Hepatitis B (HBV) and Hepatitis D (HDV) Markers According to Risk Factors and Relative Odds of Marker Positivity with 95% Confidence Intervals

		HBV		HDV	
Risk Factor	N	Positive N (%)	Adjusted* OR (95% CI)	Positive N (%)	Crude OR (95% CI)
Total	406	173 (43)	—	14 (8)**	—
Age (years)					
≤30	261	91 (35)	1.0	6 (7)	1.0
>30	144	81 (56)	1.8 (1.1–3.0)	8 (10)	1.6 (0.5–5.7)
Unknown	1	1		0	
Prison involvement (years)					
<1	166	46 (28)	1.0	1 (2)	1.0
1–4	176	81 (46)	1.8 (1.3–2.6)	8 (10)	4.9 (0.6–224)
≥5	58	42 (72)	3.3 (2.3–4.8)	5 (12)	6.1 (0.6–294)
Unknown	6	4 —	—	0	
IV drug use					
No	264	72 (27)	1.0	2 (3)	1.0
Yes	136	101 (74)	6.0 (3.6–10.2)	12 (12)	4.7 (1.0–44)
Unknown	6	0 —	—	0	
History of HBV					
No	358	137 (38)	1.0	6 (4)	1.0
Yes	17	14 (82)	5.4 (1.3–22.0)	2 (14)	3.6 (0.3–23)
Unknown	31†	22 —	—	6† —	—
Homosexual activity					
No	387	166 (43)	1.0	13 (8)	1.0
Yes	12	6 (50)	0.6 (0.1–2.2)	0 (0)	
Unknown	7	1 —	—	1	

*A multivariate model included age, years involved in prison, intravenous drug use, history of HBV, and homosexual activity.
**Denominator is number HBV positive in each risk category.
†Includes people with a history of hepatitis, type unknown, and people who refused to answer the question.

to be infected with HBV in a county jail had markers for HDV, two with antigen and two with antibody.[4] In the current study all participants were healthy at the time of screening, and the anti-delta positive results presumably represent past infection. Although one of the five cases of HBV in the outbreak that precipitated the screening/vaccination program contained IgM anti-delta, specimens from persons in this program were not evaluated for the presence of HDV antigen or IgM anti-delta. Thus some acute HDV infections could have been missed. The proportion of specimens containing anti-delta (14/173, 8 percent) from inmates in this study, however, approximates the proportion of HDV antibody noted in inmates by the study by Jacobson, *et al.*[4] As with HBV, IVDU was strongly associated with the presence of anti-delta; of the 14 inmates with markers for HDV, 12 (86 percent) gave a history of IVDU. Ponzetto, *et al*, previously reported that 10 percent of drug addicts participating in a Veterans Administration (VA) cooperative study had evidence of infection with HDV.[13]

Measures that protect against HBV infection, such as use of hepatitis B vaccine, are also effective in preventing HDV infection. Implementation of HBV preventive guidelines in corrections facilities should reduce the risk of HDV.

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Influenza Epidemics and Anencephaly

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Abstract: To explore the postulated association between maternal influenza and congenital defects of the central nervous system, 14 virologically verified epidemics in Finland, 1969–82, were studied. Mothers of 248 anencephalic children were grouped into those whose first trimester had occurred during an epidemic period and those whose pregnancy had commenced during a non-epidemic period. No significant differences in prevalence of anencephaly were noted in these groups. (*Am J Public Health* 1990; 80:473–475.)

Introduction

There is ample evidence that certain virus infections can be transmitted to the embryo during pregnancy and cause impaired fetal development expressed as congenital defects.[1] Among these potential teratogens, influenza virus has been frequently mentioned, but evidence for or against such harmful effects has remained inconclusive.[2,3] We have conducted a population study based on recorded influenza epidemics and one particular marker defect, anencephaly. This defect has several advantages for investigative purposes: its detection rate is high in stillbirths and in newborn infants, the diagnostic means have not changed during our 15-year-study period, notification rate of this severe defect is high, and finally, as it is a lethal condition, the reports to the Register of Congenital Malformations can be confirmed and completed from the compulsory death certificates. Moreover, defects of the central nervous system (CNS) have often been mentioned in the context of maternal influenza.[2-6]

Methods

The material was collected from the files of the Finnish Register of Congenital Malformations operating since 1963[7,8] and completed with data from the death certificates obtained from the Central Statistical Bureau. These sources cover the entire population of 4.9 million; the present analysis includes all mothers whose last menstruation occurred between October 1, 1968 and March 31, 1982, comprising 858,917 deliveries. Altogether 248 cases of anencephaly were recorded in children weighing 600 g or over at birth. Of these, 201 cases were single defects and, in 47, additional malformations were reported.

Influenza activity in the Finnish population has been monitored by the Central Public Health Laboratory since 1969 with virus isolation and serodiagnostics of acute respiratory infections. Our study covered 14 epidemic periods during which influenza etiology was confirmed. Excluding sporadic cases and small local outbreaks, the epidemic time

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TABLE 1—Occurrence of influenza epidemics in Finland and serological infection percentages in samples of serum pairs from pregnant women.

Epidemic periods	Number of serum pairs (a)	Influenza A H1N1 (%)	Influenza A H3N2 (%)	Influenza B (%)
January 69–February 69		–	(b)	–
December 69–February 70		–	(b)	–
April 71–May 71		–	–	(b)
November 71–December 71	137	–	27	–
January 73–February 73	91	–	20	–
January 74–April 74	106	–	23	18
January 75–March 75	144	–	17	–
January 76–March 76	397	–	15	–
March 77–April 77	636	–	14	7
January 78–March 78	651	1	8	–
January 79–April 79	549	3	3	2
January 80–February 80	576	–	25	–
January 81–March 81	674	10	2	4
February 82–June 82	291	1	1	1

(a) The pairs consisted of pre-epidemic and postepidemic specimens

(b) Epidemic activity recorded, but infection frequency among pregnant women not studied

was 37 months of the total 162 months monitored during the study period. To create a "non-epidemic" control material, the month preceding and following each epidemic was excluded because of timing uncertainties. Altogether, 26 such borderline months with 46 anencephalic children were thus excluded.

Since the end of 1971, the influenza surveillance program has included annual monitoring of serological infection frequencies among pregnant women in Finland.[9] Serum pairs of pre-epidemic and post-epidemic specimens from a total of 4,252 women were studied for haemagglutination-inhibiting antibodies to influenza A and B viruses including the epidemic variant, and, for detection of anamnestic antibody response, the serotypes of the near past (Table 1). A fourfold or higher rise in antibody titers was considered indicative of serological influenza infection.[9]

For the occurrence of anencephaly in infants, mothers whose first trimester (14 weeks from the start of the mother's previous menstruation) had coincided wholly or in part with a virologically verified influenza epidemic (407,200 mothers) and mothers whose first trimester had occurred completely during a non-epidemic period (314,841 mothers) were compared after exclusion of the borderline months. Power calculations were performed according to Schlesselman.[10] The heterogeneity test was performed and confidence limits for rate ratio estimates were calculated as suggested by Miettinen.[11] Trend analysis for the proportion of rates of anencephaly was performed according to Armitage.[12]

Results

Epidemics caused by influenza A and B viruses, and serological infection frequencies in the follow-up samples of pregnant women are listed in Table 1. Figure 1 shows the rates of anencephaly when the first trimester had occurred during the influenza epidemics or during the non-epidemic periods. In the epidemic group, 117 anencephalics were reported and in the non-epidemic group, 85; prevalence 0.29 per mil and 0.27 per mil, respectively.

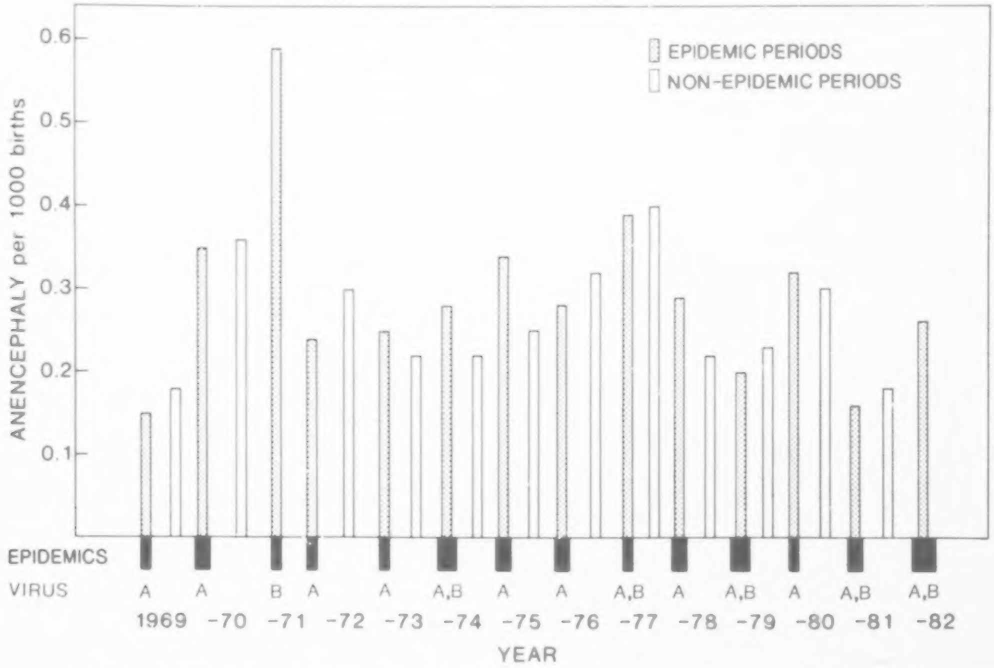

FIGURE 1—Prevalence of Anencephaly in Children Whose first 12 Weeks of Development Coincided either with an Influenza Epidemic or with the Non-Epidemic Periods

When the rates of anencephaly were examined in relation to the serologically verified influenza epidemics, the April-May 1971 B-influenza epidemic showed an apparently high rate in comparison with other periods (Figure 1). However, the heterogeneity test indicated that within the study period of 1969–82, none of the observed rates showed statistically significant elevation ($p = 0.75$). It was evident that influenza epidemics in general were not associated with the occurring of anencephaly. The rate ratio point estimate for anencephaly was 1.0 (95% confidence limits 0.8, 1.3), when the infants of mothers whose first trimester had coincided with an influenza epidemic were compared with the infants whose sensitive period of development had occurred during a non-epidemic period. Trend analysis showed no change in the rate of anencephaly for the period of 1969–82 (regression slope −0.0000028, SE 0.0000026).

Discussion

Two previous studies in the same population suggest, as we do here, that an influenza virus infection of the mother might not be a major cause of neural tube defects: serological screening of mothers to defective children showed no elevated titers against influenza virus antibodies as compared with matched-pair controls,[13] and CNS defects showed no seasonal variations in 1965–73 despite the definite winter excess of influenza.[14] The latter finding confirms similar observations in other populations.[15]

An exception to the negative findings listed above seems to be provided by the Asian influenza pandemic in 1957 caused by a new subtype (H2N2) of the influenza A virus. Several independent studies reported an association of this epidemic and an increased risk of defects of the CNS, anencephaly included.[14,16–20] Caused by a new subtype of the virus, the epidemic attacked a highly susceptible population, which might explain the association. In the present material, a similar effect could not be found for the 1968/69 pandemic in which the new subtype H3N2 was involved. The present analysis did not reveal any association between an increased risk of anencephaly and maternal influenza infection. The power of the study was about 80 percent to detect a rate ratio of 1.4 for all epidemics combined and almost 100 percent to detect rate ratios of 1.7 or greater ($\alpha = 0.05$, one-sided).

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Rabies Vaccine Adsorbed: Neutralizing Antibody Titers after Three-Dose Pre-exposure Vaccination

BYRON S. BERLIN, MD

Abstract: Field trials at several schools of veterinary medicine showed that three-dose pre-exposure rabies vaccination with Rabies Vaccine Adsorbed developed by the Michigan Department of Health elicited neutralization antibody in practically all recipients two to three weeks after immunization. Titers declined during the first six months after vaccination. However, by 18 to 24 months, 98 percent of recipients still had titers equal or greater than a 1:5 dilution of serum. (*Am J Public Health* 1990; 80:476–478.)

Introduction

Initial clinical trials with Rabies Vaccine Adsorbed (RVA), a new tissue culture derived rabies vaccine developed by the Michigan Department of Public Health,[1] were done with a two-dose pre-exposure immunization schedule.[2] The studies showed that this new vaccine induced satisfactory antibody titers two to three weeks after vaccination and prepared recipients for a vigorous response to booster vaccination. Additional clinical trials, summarized in the present study, were conducted to ascertain antibody responses after administration of this vaccine by pre-exposure immunization schedules recommended by the Immunization Practices Advisory Committee (ACIP), Centers for Disease Control, i.e., three doses of rabies vaccine of tissue culture origin, one dose each day on days 0, 7, and 21 or 28 days.[3]

Methods

Vaccine

The vaccine used in the present study was previously described.[1] Briefly, RVA contains non-infectious, B-propiolactone-treated rabies virus which is adsorbed onto aluminum phosphate. The virus was grown on diploid fetal rhesus monkey lung fibroblasts (FRh1-2), the rhesus analogue of the WI-38 cell. The antigenic potency of each lot of RVA exceeds 2.5 International Units (IU).

Subjects

The study consisted of 24 separate clinical subtrials with a combined enrollment of 1,735 students and staffs from five schools of veterinary medicine: 1,647 with no prior vaccination against rabies and 88 who had not responded to Duck Embryo Rabies Vaccine. The number of volunteers in individual subtrials varied from 16 to 127 and the median size was between 68 and 74 persons. Each person received three intramuscular injections of RVA, 1.0 mL per dose, at 0, 7, and 21 or 28 days and then was bled two weeks after the final dose of vaccine and, depending on the subtrial in which they were included, at six months and nine to 14 months. Volunteers remaining on the campuses were bled 18 to 24 months after vaccination (Table 1). Serum samples were sent to the Michigan Department of Public Health laboratories for antibody determination. Each volunteer provided a brief medical history and signed an informed consent before vaccination.

Serology

Serum neutralization titers were ascertained by the Rapid Fluorescent Focus Inhibition Test (RFFIT)[4] beginning with a 1:5 dilution of the serum and ending with a 1:625 dilution using five-fold dilutions. RFFIT titers, given as International Units (IU), were computed by dividing the reciprocal of the dilution titer of the subject's serum by the reciprocal of the dilution titer of a standard serum, designated as 1.0 IU, tested at the same time. RFFIT titers less than 1:5 dilution of serum, approximately 0.1 IU were considered below the minimally acceptable titer recommended by the ACIP.[3] Titers from 0.2 to 0.5 IU were classed as acceptable because they fell between 0.1 IU and the minimally acceptable initial RFFIT response to post-exposure immunization recommended by the World Health Organization (0.5 IU).[5] Titers above 25 IU were grouped together because the maximum dilution for testing the sera was 1:625.

Results

Ninety-six percent (1,673/1,735) of the vaccinees were bled two weeks after the last dose of the vaccine series (Table 1). Nineteen (1 percent) had either withdrawn voluntarily or were dismissed from the study because of illness or reaction to the vaccine. Forty-three (2 percent) failed to return for bleeding. All but one of the 1,673 who were bled at two weeks had rabies neutralizing antibody titers greater than 0.1 IU.

Subsequently, 91 percent (419/461) of volunteers initially enrolled in five of the 23 subtrials were bled six months after the vaccine series (Table 1) and, of these, 415 (99 percent) had RFFIT titers greater than 0.1 IU. Nine to 14 months after completing the vaccine series, blood samples were obtained from 761/984 (77 percent) volunteers enrolled in 13 subtrials. Of these, 748 (98%) also had RFFIT titers equal or greater than 0.1 IU. Finally, 308 persons (52 percent of the initial

TABLE 1—Rabies Vaccine Study Plan

		Time and Number of Persons Bled			
Number of Subtrials	Number of Volunteers	Two Weeks	6 mos	9–14 mos	18–24 mos
7	360	348			
1	101	99	84		
3	251	249	234	179	
6	431	413	N.D.*	301	
1	109	109	101	N.D.	98
4	366	350	N.D.	281	133
2	117	105	N.D.	N.D.	77
Total (24)	(1,735)	(1,673)	(419)	(761)	(308)
Percent 0.1 IU		1	4	13	5
Number of initial volunteers		1,735	461	984	592
Percent of initial volunteers bled		96	91	77	52

*Not Done.

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TABLE 2—Persistence of RFFIT Titers after Pre-exposure Vaccination with Rhesus Diploid Rabies Vaccine

Number Enrolled	Antigen content*	Schedule Days**	Median Titer***	Bleeding Time	RFFIT Titer Range***							
					<0.2	0.2–0.5	0.6–1.0	1.1–5.0	5.1–15.0	15.1–24.9	>25	Number
109	2.6	0-7-28	6.4	2 wks	0	0	0	25	52	12	5	94
			2.2	6 mos	0	2	23	57	9	2	1	94
			1.2	18 mos	1	23	22	46	2	0	0	94
74	2.6	0-7-21	5.4	2 wks	0	0	1	9	33	9	3	55
			0.3	22 mos	2	26	13	8	5	1	0	55
88	3.0(a)	0-7-28	5.4	2 wks	0	0	1	3	21	7	2	34
			0.4	2 yrs	7	12	10	3	1	0	1	34

*Antigenic Potency expressed as International Units (IU).
**Days of Injection.
***RFFIT Titer expressed as International Units (IU).
(a)Vaccine contained 1:10,000 thiomerosal as a preservative.

volunteers in seven subtrials) volunteered to be bled from 18 to 24 months after vaccination. Ninety-eight percent (303/308) still had titers equal to or greater than 0.1 IU.

The findings in three representative clinical subtrials in which neutralizing antibody titers are presented only for persons from whom successive blood samples were obtained at each indicated time are shown in Table 2. Successive blood samples were obtained from 86 percent (94/109), 74 percent (55/74), and 39 percent (34/88) of volunteers initially enrolled in the respective trials. Two weeks after the third and final injection of RVA, median RFFIT titers in the three subtrials were 6.4, 5.4 and 5.4 IU, respectively. All recipients in these subtrials had neutralizing antibody at two weeks and, in those tested, at six months after vaccination. However, neutralizing antibody titers six months after vaccination were lower than those at two weeks. Eighteen months to two years post-vaccination, median RFFIT titers were from to 0.3 to 1.2 IU and 261/271 (96 percent) of the volunteers had titers equal to or greater than 0.1 IU.

Discussion

The findings summarized in the present study show that the seroconversion rate after pre-exposure vaccination with RVA, given by the currently recommended three dose schedules, approximated 100 percent. However, the duration of the antibody response after initial pre-exposure rabies immunization is as important a concern as the seroconversion rate in evaluating immune responses to rabies vaccines because neutralizing antibody, actively induced by vaccination or passively acquired, is a key factor in protection against rabies infection.[6–9] Yet, determination of the duration of responses to vaccination in sizable populations is complicated by reason of attrition of volunteers. Attrition rates at six months and at nine to 14 months in this study were 9 percent and 23 percent, respectively; these rates probably have little effect on the interpretation of the results in view of the finding that 94 percent of recipients or more had titers equal to or greater than 0.1 IU. On the other hand, even though the attrition rate for those bled 18 months to two years after vaccination was 48 percent, the percent of volunteers showing neutralizing antibody at that time is considered a good approximation of persistence of antibody because titers at two weeks in those who were bled later did not appear different than those in persons who were not bled at the later time.

Although RVA and human diploid cell culture vaccine (HDCV) have not been compared to each other in concurrent clinical trials, the seroconversion rate, persistence of rabies neutralizing antibody, and adverse reactions after pre-exposure immunization with RVA appear practically the same as those obtained with HDCV.[2,10–19]

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Health Effects of a Thorium Waste Disposal Site

G. Reza Najem, MD, PhD, and Lisa K. Voyce, MSENE

Abstract: A case-control study of 112 households residing in the vicinity of a thorium waste disposal site found a higher prevalence of birth defects (RR 2.1) and liver diseases (RR 2.3) among exposed than the unexposed group. The numbers were quite small and the confidence intervals wide, however, so that no definite conclusions can be drawn from these data. (*Am J Public Health* 1990; 80:478–480.)

Introduction

Several studies have appeared in the medical literature analyzing possible adverse health effects for persons living in the vicinity of hazardous waste disposal sites.[1-3] To our knowledge there is no published study in which the health effects of a thorium waste disposal site (TWDS) on the general population have been investigated, except for a community-based study in China.[4] That study found no deleterious health effects from exposure to thorium, but had very low statistical power.

The main concern about the possible biological effects of thorium (Th) with a half life of 1.4×10^{10} years is based on its radioactive characteristics and isotopes.[5] Thorium^{232} is the parent of various radio-isotopes including ^{228}Ra (radium), ^{228}Ac (actinium), ^{228}Th, ^{224}Ra, and ^{220}Rn (radon)[5]. Current interest is based in part on increased use of thorium as an energy source, e.g., in reactors using ^{232}Th to produce ^{233}U.[6]

Industrial and mining exposure to thorium have resulted in an increased incidence of lung cancer, pancreatic cancer, colorectal cancers, chronic respiratory diseases, liver damage, and other serious illnesses.[7-9] Skeletal sarcomas, thorotrastomas, and sarcomas of the reticuloendo-thelial system have been reported from a study of patients receiving intravenous injections of thorotrast, a collodial solution of thorium oxide that was used as a contrast medium for diagnostic radiographic studies from the 1930s to the 1950s.[10]

The purpose of this study was to conduct a survey of health effects of a thorium waste disposal site among members of the general population of Wayne, New Jersey, USA.

Thorium Waste Disposal Site

The site is located on a 6.5 acre tract of land in Wayne Township, Passaic County in northern New Jersey, USA.* The population density in Wayne is 1,867.2 per square miles and annual per capita income is $7,725.[11] The thorium waste burial site lies within 100 feet of the sidewalk. Commercial businesses, a nursing home, a school bus maintenance yard, and residential homes are adjacent to the site. A drainage ditch lies at the northern and eastern borders, and drains into a brook at the northeast corner of the burial field. The brook (a tributary to the Pompton River) flows through residential areas (copy of map available on request). Fresh water wetlands lie approximately 1,800 feet to the west of the thorium waste disposal site.* From 1948 until 1971, a company extracted thorium and rare earths from monazite ore, for usage by the Atomic Energy Commission.* During its active years, the plant waste was dumped in backyard sludge piles. The wastes have been buried only in the last few years, after spending over 20 years above the ground.**

Surveys of the New Jersey Department of Environmental Protection showed readings of 40-1491 uR/hour (uR = micro-rad, micro being 10^{-6}) in various regions classified as contaminated areas; Sheffield Brook contamination occurs within 10 meters of both sides. Soil samples revealed Thorium (^{232}Th), Radium (^{238}Ra) and Uranium (^{238}Ur) in the contaminated regions. Water samples from Sheffield Brook and the drainage ditch showed mainly alpha and beta activity.

Most of the air concentrations (ambient and indoor) were within the background radiation levels (approximately 150 mrem/year) The US Environmental Protection Agency (EPA) standard for home concentration is below 4 pCi/liter.***

The residents of the households in the vicinity of the thorium waste disposal site and in the vicinity of contaminated Sheffield Brook have not been tested for radioactive materials uptake.

Methods

The exposed population included 76 households within three blocks adjoining the contaminated areas. The 76 comparison (unexposed) households included all houses located on about 9th and 10th blocks distance from the thorium waste disposal site and contaminated Sheffield Brook. Households were accepted for inclusion in the study if they had resided at the present address for at least five years.

Eleven and 10 households of the exposed and unexposed groups, respectively, did not meet the criteria for eligibility, 10 and nine households, respectively, did not participate in the study due primarily to members not being at home at the time of scheduled interviews or having moved out without leaving forwarding addresses.

Face-to-face interviews of all 112 households (362 members), using a structured questionnaire, were carried out by one interviewer who did not know exposure status; 32 household members who did not agree to a face-to-face

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*New Jersey Department of Environmental Protection: Site Descriptions for Hazardous Waste Sites in New Jersey. Division of Waste Management, Hazardous Site Mitigation Administration, 1984/85.

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**New Jersey Department of Environmental Protection. Unpublished data.

***US Environmental Protection Agency, Office of Radiation Programs, Washington, DC 20406.

Further details on measurements of exposed areas are available on request to author (Najem).

TABLE 1—Demographic Variables among Exposed and Comparison Groups

Variables	Exposed Group	Comparison Group
# of Persons studied	171	191
# of Households studied	55	57
% of Persons Age 50 & over	29	32
% of Persons Age 18 & under	27	26
Duration of Residence		
Mean duration*	13.3 Years	17.2 Years
% Under 10 Years	42	23
% 10–20 Years	23	9
% Above 20 Years	35	68
Annual Family Income		
% Less than $15,000	9	9
% $15,000–$49,000	65	63
% $50,000–$59,000	20	21
% $60,000+	6	7

*Mean duration difference = 3.9 years (95% CI = −.71, −7.23)

interview were interviewed by telephone. For each disease or condition reported by the participants, we asked respondents whether they obtained the diagnostic information directly from their physicians; we did not obtain or review medical records.

The risk ratios associated with the disease were estimated by calculating relative risk ratios.[12] Cornfield's 95% confidence intervals (CI) and Miettinen's 95% CI were computed by using True Epistat computer software.[13] The Mantel-Haenszel method of controlling for confounding variables was used by using the SAS (Statistical Analysis System) computer software.[14]

Results and Discussion

There were no important differences in age, sex, race, income, marital status, or religious affiliation of the exposed and comparison groups, but the comparison group had resided longer in the community (Table 1). There were no differences in smoking history, alcohol consumption, frequency of eating local produce, and history of working in a chemical industry, or in reported history of cancer between the two groups (Table 2).

Although there was a higher prevalence of birth defects and liver diseases among the exposed than the control group, numbers are small and confidence intervals wide (Table 3). After adjusting for duration of residency and alcohol consumption, the estimated risk ratios were very similar, except for 1.1 (95% CI = .05, 78.4) for low birth weight and 2.9 (95% CI = .12, 168.24) for liver diseases.

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This study was presented in part at the XI Scientific Meeting of the International Epidemiological Association, Helsinki, Finland, August 1987.

TABLE 2—Comparison of Risk Factors among Exposed and Comparison Individuals and Households

Individuals	Exposed Group (N: 171)	Comparison Group (N: 191)	Risk Ratio (95% CI)
Ever smoked	57 persons	64 persons	0.94 (.63, 1.57)
Ever smoked cigarettes	96% of smokers	94% of smokers	
Average # cigarettes smoked per day	21 cigarettes	18 cigarettes	
Average # of years smoked	19 years	22 years	
3+ alcoholic drinks per day at present	10 persons	12 persons	0.93 (.36, 2.37)
3+ alcoholic drinks per day in past	37 persons	32 persons	1.37 (.78, 2.40)
Drinking 5 oz per week at present	13 persons	25 persons	0.54 (.25, 1.15)
History of working in a chemical industry	20 persons	31 persons	0.7 (.36, 1.30)
Households			
Grow vegetables	19 households	23 households	
Cancer among first degree relatives	31 households	31 households	1.1 (.48, 2.44)

TABLE 3—Diseases and Adverse Birth Outcomes among Exposed and Comparison Groups

	Exposed Group N = 171	Comparison Group N = 191	Risk Ratio	95% CI
Low Birthweight	3	6	0.6	.11, 2.52
Skin diseases	7	13	0.6	.21, 1.62
Anemia	6	7	1.0	.27, 3.24
Spontaneous abortion	13	14	1.0	.44, 2.43
Liver diseases*	4	2	2.3	.35, 12.0
Birth defect**	9	5	2.1	.62, 7.25
Cancer***	11	10	1.2	.48, 3.26

*Including: jaundice and hepatomegaly.

**Born with birth defects at the present address including: cerebral palsy, congenital heart disease, inguinal hernia, hemangioma, foot deformities, other bone and joint deformities.

***Including: leukemia, bone cancer, lung cancer, larynx cancer, and nose cancer.

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Comparison of Antineoplastic Drug Handling Policies of Hospitals with OSHA Guidelines: A Pilot Study

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Abstract: Hospital antineoplastic drug handling policies of 24 hospitals in eight Southwestern Ohio counties were compared with recommendations of the 1986 Occupational Safety and Health Administration (OSHA) guidelines. Although most study facilities where antineoplastics are handled have policies, content varies and is generally less complete than OSHA guidelines, particularly regarding storing, transporting, and disposing of drugs; managing equipment; informing personnel of risk; and surveillance. Recommendations for personal protection concur more closely with OSHA guidelines than do other content areas. (*Am J Public Health* 1990;80:480–481.)

Introduction

Evidence of probable risk associated with staff handling of antineoplastic drugs led the US Occupational Safety and Health Administration (OSHA) to issue comprehensive guidelines covering staff training and surveillance; use and care of protective equipment, apparel, and techniques; and storage, disposal, and transport of drugs and contaminated equipment.[1] Concerned about moral and legal issues, many hospitals have also developed institutional policies.

This study investigated the status of facility policies for handling antineoplastic drugs in hospitals and home health care agencies in eight southwestern Ohio counties and compared policy content to OSHA guidelines.

Methods

In July 1986, a questionnaire, modeled after Tortorici's[2] was mailed to the directors of the 32 hospitals and 25 community health agencies in the eight predominately rural counties comprising Region I of the Cancer Control Consortium of Ohio. Respondents were requested to enclose their facility's antineoplastic drug handling policy. Policy content was compared with that of the 1986 OSHA guidelines using an investigator-designed tool for systematically coding content relative to OSHA guidelines. For each item, the facility policy is rated 0 if the item is absent and 1 if the item is included. The instrument yields a maximum total score of 219 and five subscores: 1) General Guidelines (37); 2) Protective Equipment (14); 3) Facilities for Mixing and Storing Agents (15); 4) Protective Apparel (49); and 5) Protective Techniques (104). The content analysis was conducted by one investigator (KD).

Results

Twenty-four hospitals and 11 community health agencies responded, a response rate of 75 percent for hospitals, 44 percent for home health agencies. Administrators of 20 of the hospitals report antineoplastics are mixed and administered at their facility and one home health agency administers premixed antineoplastics; 18 of the 20 hospitals and the home health agency have implemented an antineoplastic drug handling policy. Eleven hospitals submitted their policy for analysis.

All 11 policies are less comprehensive than OSHA guidelines. Total policy scores of individual facilities ranged from 26 to 92 of a possible 219. The mean total score was 50.6 (median score 41). Areas least addressed under "General Guidelines" were maintaining files on specific antineoplastic agents and providing staff with information on known risks, medical policies, safe handling of investigational drugs as they are introduced, and follow-up evaluation of staff knowledge and competence regarding antineoplastic drug handling. Least addressed under "Protective Equipment" were procedures for installation, maintenance, and decontamination of hoods, while unaddressed content under "Facilities for Mixing and Storing Agents" related to reserving a hood only for mixing antineoplastics, posting warning signs, listing drugs covered by cytotoxic policies, and labeling stored antineoplastics. Under "Protective Apparel" recommendations for use of protective apparel during drug mixing conformed better to OSHA guidelines than recommendations for administration, while those for excreta handling and cleaning up spills were generally inadequate. Recommendations for glove and gown use were more often consistent with OSHA recommendations than for use of masks or eye protection. Under "Protective Techniques" least frequently

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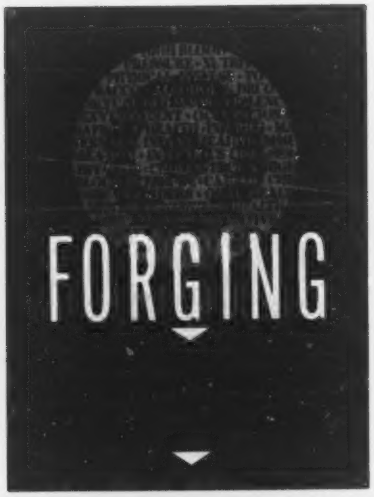

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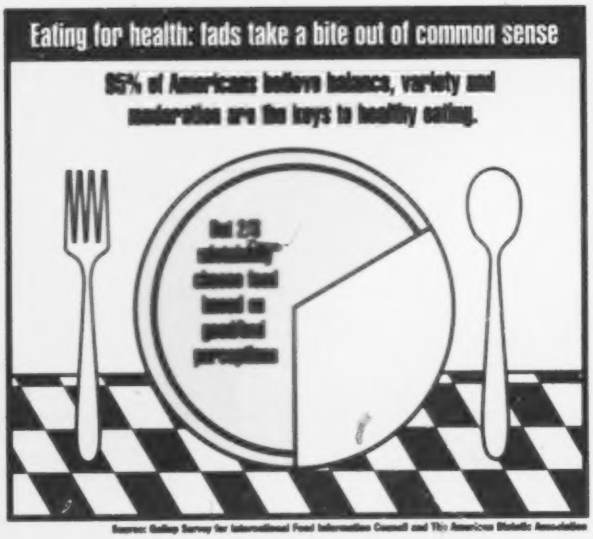

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Improving Influenza Vaccination Performance in an HMO Setting: The Use of Computer-Generated Reminders and Peer Comparison Feedback

MARY B. BARTON, BA, AND STEPHEN C. SCHOENBAUM, MD, MPH

Abstract: We evaluated a program for improving influenza immunization performance in a health maintenance organization (HMO). The HMO implemented several interventions successively from 1984–87: a postcard reminder to members at high risk for complications of influenza, a computer-generated reminder to the physician at the time of any primary care visit by high-risk patients, performance feedback to chiefs of service, and, finally, retrospective feedback to each physician comparing his/her performance with that of the other physicians. We examined immunization rates for a group of members older than age 65, a high-risk group under age 65, and a group of diabetic members who had not been subject to the reminders (vs a group who had been covered by the program).

Vaccination rates were increased in those diabetic members who received reminders. Nevertheless, among members younger and older than age 65 whose experience was observed over three flu seasons, a significant increase in vaccination rates was not achieved until physician feedback was added to the program. We conclude that each element of the reminder and feedback program has contributed to the overall increase in vaccination rates at the HMO and that effective ongoing influenza immunization programs can be implemented in practice settings with appropriate systems support. (*Am J Public Health* 1990; 80:534–536.)

Introduction

Influenza accounted for more than 10,000 excess deaths in the United States in 19 different epidemics between 1957 and 1986.[1] Yet, an influenza vaccine containing a strain similar to that causing epidemic disease has been found to be up to 80 percent effective at protecting the elderly, chronically ill population against hospitalization and death from acute respiratory disease during an influenza epidemic.[2] With the goal of reducing excess morbidity and mortality from influenza, the Immunization Practices Advisory Committee (ACIP) of the Centers for Disease Control recommends the annual vaccination of specific groups with chronic medical conditions who are at high risk for complications of influenza. In the early 1980s, the US Public Health Service's stated goal was vaccination of 80 percent of those at highest risk for morbidity and mortality from influenza infection. Unfortunately, estimates indicate that fewer than 25 percent of high-risk individuals were receiving an influenza immunization each year.[3]

Despite the good intentions of their physicians, Americans do not necessarily receive the benefits of available preventive health measures such as cancer screening and immunizations.[4] Several randomized controlled trials have shown that systems for physicians, such as reminders and feedback, can improve delivery of preventive measures.[5–7] In addition, postcard reminders to patients have been shown to increase influenza vaccination among elderly populations, including those in a health maintenance organization (HMO) setting.[8–10] Although these trials demonstrate the efficacy of practice support systems, there is little evidence that reminders and feedback actually have been adopted in practice settings or that they are effective in an ongoing practice-based program as distinct from a special study. The Health Centers Division of Harvard Community Health Plan (HCHP) is a large staff model HMO in the Boston area. Most of HCHP's health centers use a computerized medical record system for documenting ambulatory care.[10] This system has also been used as a tool in controlled trials of limited size to monitor clinician performance and provide feedback.[7,12]

In recent years HCHP has instituted a program to improve influenza vaccination of its high-risk members. The program includes educating members about the need for and availability of flu vaccine and consists of a series of systems for member reminders, physician reminders, and physician feedback on performance. The program has developed in increments. The HMO has not considered it to be research but rather part of its efforts to improve clinical operations and quality of care. This paper presents a description of the program as it has evolved, an evaluation of HCHP's influenza immunization performance and an assessment of the effectiveness of the reminder and feedback systems.

Methods

The membership of the Health Centers Division of HCHP grew from 178,000 in 1984–85 (Flu Season #1 in this evaluation) to 227,000 in 1986–87 (Flu Season #3). During that time, an average of approximately 100 internists practiced adult primary care at seven health centers in the metropolitan Boston area.

The first year's intervention involved postcard reminders to HMO members, educational materials (e.g., posters for waiting areas and examination rooms, pharmacy bag notices), and a reminder message to clinicians at the front of the summary of the computerized record which was prepared for each scheduled primary care visit for a high-risk patient under age 65. This was supplemented in the second year by chart reminders for persons over age 65 and feedback of performance to service chiefs. In the third year these actions were further supplemented by feedback of performance to individual physicians and by the periodic distribution to physicians of lists of patients not yet immunized.

In 1983–84, the year before the influenza vaccination program began, a survey of records of members age 65 and over revealed immunization rates of 24 percent in centers using the automated record and 22 percent in the one health center using a manual record. In order to detect change in vaccination rates over the three succeeding flu seasons, we selected a cohort of members representing the high-risk group under age 65 and a cohort of members age 65 and over.

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The latter were chosen from among those members who had belonged to HCHP for four or more years as of June 1987, had turned 65 by September 1984, and had designated a specific physician as their primary provider. Lists of members meeting these criteria were obtained for each health center and stratified by provider; a random sample was then selected from each center's list. By using computerized enrollment files it was possible to generate a similar list for the center not using the automated medical record.

The cohort of persons under age 65 at high risk for complications for influenza was identified from the computer-generated list used to send out reminder postcards in early autumn and to flag charts with a reminder message at the time of flu season visits. The computer algorithm used to create this list consists of a search for certain diagnostic codes corresponding to high-risk categories which were entered during a specific window of time in the prior year (roughly October 1-June 1 each year).

In order to assess the specific effect of a reminder, we focused on a single diagnostic group—persons with insulin-dependent diabetes: HCHP maintains a Plan-wide registry of insulin-dependent diabetics culled from diagnostic codes entered in the computerized medical records and also from all pharmacy slips recording prescriptions for insulin. In comparing this list with the flu vaccine reminder list for Flu Season #3, some members from centers with the automated record were identified who were not on the reminder list despite their having a high-risk diagnosis. Reasons for diabetic members not appearing on the flu vaccine reminder list include: joining HCHP after the flu vaccine reminder list had been made for the year, having no visit to HCHP during the window of time that computer records are searched for high-risk diagnoses, transferring care to another health center and thus a different member identification number, or having a very recent initial diagnosis of insulin-dependent diabetes mellitus (i.e., after the search window for creating the reminder lists).

Because the mean age of the non-reminder group (40.8 years) was significantly lower than that of the overall group of diabetics under age 65 on the vaccine reminder list (48.3 years), we used the subset of both groups between the ages of 40 and 65 for comparison of vaccination rates (respective mean ages ± SEM: 53.4 years ± .7 and 54.5 years ± .5).

Records of members at the computerized centers were searched by computer for the therapy code for influenza vaccine. Records for members at the non-automated center were reviewed by hand.

Data analysis was performed using SAS version 5.

Results

The overall vaccination rates at the computerized centers for the three flu seasons starting in 1984 are shown in Table 1. Vaccination rates increased substantially in both cohorts in Flu Season #3 (1986–87), which corresponded to the institution of the physician feedback mechanism. Immunization rates at the center without the automated record remained stable at 24–33 percent throughout the three flu seasons.

The focused study of diabetic members revealed that those on the reminder list were both more likely to have a visit during flu season and more likely to be vaccinated if they had a visit to their primary clinician (Table 2). The difference in visit rates probably reflects the effect of the postcard reminder on encouraging flu season visits. It is important to recall, however, that the reasons that prevented the "no reminder" group from being included on the computer's flu vaccine reminder list are directly related to lower visit rates in the previous year and these reasons may contribute to the continued lower visit rate in this group. In contrast, the difference in vaccination rates for those with visits—73 percent for those receiving the reminder versus 42 percent for those not getting the reminder—ought to reflect the effect of the computer-generated chart reminder on the physician.

TABLE 1—Influenza Immunization Rates in an HMO during Flu Seasons #1–3, 1984–1987

	Over Age 65 (n = 647)		High Risk, Under 65 (n = 198)	
Flu Season	% Immunized	Annual Change in Rate*	% Immunized	Annual Change in Rate*
#1	42		38	
		0 (−5, 5)		3 (−7,13)
#2	42		41	
		22 (17,27)		14 (4, 24)
#3	60		55	

*Difference between rates in successive seasons (95% confidence limits)

TABLE 2—Vaccination Status of Diabetic HMO Members, Ages 40–65

Vaccination Status	Reminder (%)	No Reminder (%)	Difference*
Vaccinated	55.9	27.0	28 (16, 40)
Patient refusal	0.0	1.8	−2 (−4, 0)
Not indicated	4.9	1.8	3 (−1, 7)
No flu season visit	23.8	33.3	−10 (−21, 1)
Not vaccinated despite a flu season visit	15.4	35.1	−20 (−33, −7)
Total in group	143 (100.0%)	111 (100.0%)	

*Difference in percentage between reminder and no reminder groups (95% confidence limits)

Discussion

As noted at the outset, there have been several controlled studies of various interventions to improve performance of periodic health measures. All have shown efficacy of the specific intervention under study. None, however, has been a demonstration of sustained program effectiveness; nor has any of the studies attempted to determine the effect of additive interventions. The present report demonstrates that it is possible to achieve improvement in annual influenza vaccination performance in a practice setting using a combination of reminder and feedback techniques.

At least one other HMO has reported achieving performance similar to HCHP's but with a different combination of techniques. Group Health Cooperative of Puget Sound (GHC), which has a large elderly population, has reported vaccination rates for its members over age 65 of 48 and 53 percent in 1986–87 and 1987–88, respectively.[13] GHC sends reminder letters to its members over age 65, since they can be identified from automated enrollment files. The program runs mass vaccination clinics over a period of several weeks. Volunteer senior citizens help staff the clinics. The program, however, does not specifically target younger persons at high risk of complications of influenza.

HMOs can increase the cost-effectiveness of influenza vaccine delivery both by running mass immunization clinics and by appending the immunization process onto an encounter which was precipitated for another reason. In either case the cost per immunization delivered should be less than if a specific visit has to be set aside solely for vaccination.

The routine data on the flu program generated from HCHP's annual feedback show that in the centers with the computerized medical record approximately 70 percent of those members on the reminder list who visit their doctor during the flu season actually get the vaccine; 15 percent refuse; and for 15 percent there is no documentation of whether they were offered vaccine, got it, or refused. These are very positive results: many more HCHP members thought to be at greatest risk of complications from influenza, persons with chronic cardio-pulmonary disease under the active care of physicians, get vaccinated than national[3] or local[14] statistics indicate occurs elsewhere. Since physicians appear to be able to give the vaccine at HCHP and patients appear to be willing to accept it most of the time it is offered, neither physicians nor patients seem to be major barriers to more effective vaccination.

As a result of the evaluation, HCHP has continued the influenza immunization program, including: postcard and chart reminders for all members over age 65 and high-risk members under age 65, periodic distribution of lists of unimmunized members to clinicians during the vaccination season, and feedback of performance statistics to individual clinicians and chiefs of service. Since the basic computer programming of the reminders has been completed and the computer generates the lists and statistics for feedback, this influenza immunization program requires a minimum of effort to sustain.

ACKNOWLEDGMENTS

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ERRATUM

In: Marin G, Vanoss B, Perez-Stable EJ: Feasibility of a telephone survey to study a minority community: Hispanics in San Francisco. Am J Public Health 1990;80:323–326.

The name of the second listed author was incomplete as published. It should have read: Barbara Vanoss Marin. We regret that her last name was omitted on the table of contents cover and on the first page of the article.

Tobacco Excise and Declining Tobacco Consumption: The Case of Papua New Guinea

SIMON CHAPMAN, PHD, AND JEFF RICHARDSON, PHD

Abstract: Excise elasticities of demand for non-cigarette tobacco and cigarettes were calculated for Papua New Guinea for the 14 years 1973–86. Respectively, these were −0.50 and −0.71. This means that 10 percent increases in cigarette and non-cigarette tobacco excise are associated with an estimated 7.1 percent fall in demand for cigarettes and a 5 percent decrease in the demand for tobacco. As tax (excise) elasticities would normally be significantly smaller than price elasticities, these results imply that the consumption of cigarettes and tobacco in Papua New Guinea are more responsive to prices than in the United States and other Western countries. The level of excise is therefore an important and practical instrument for the control of consumption. These elasticities appear to be the first reported for a developing country. It is suggested that if they are indicative of consumer behavior in lower income countries then increasing tobacco and cigarette excise should be considered as an important strategy for the control of smoking in these countries which, because of their large populations, are huge markets for tobacco products and thus major targets for tobacco control measures. (*Am J Public Health* 1990; 80:537–540.)

Introduction

An expert working group convened by the American Heart Association recently recommended an increase in the rate of excise taxation on tobacco products and the resulting increase in tobacco prices as the leading priority for the United States among a range of smoking control strategies.[1] The success of the strategy would be directly proportional to the elasticity of demand which is defined as the percentage change in quantity demanded divided by the percentage change in price. Recent studies indicate that the price elasticity of demand for cigarettes by adults in the USA is approximately −0.4, meaning that a 10 percent increase in price should create a 4 percent decrease in demand.[2] A substantial portion of this decrease seems to result from a decrease in smoking prevalence through cessation and falling incidence, rather than through less smoking by continuing smokers.[2]

The elasticity of demand for cigarettes has been shown to be considerably greater in groups with lower levels of disposable income, namely teenagers (−1.4)[3] and people of low socioeconomic status[4]; such groups will stop smoking at a faster rate than those with higher disposable incomes if the price of cigarettes rises, and they may be influenced to not start smoking because of the additional financial burden. However, no studies appear to have been published on the effects of price changes on demand in developing countries, where disposable incomes are typically very low. Despite this, per capita consumption of cigarettes and smoking prevalence in some places is rising at alarming rates,[5] and the promotional activities of the tobacco industry are generally unchecked by any government regulations.[6]

Smoking in Papua New Guinea

Papua New Guinea (population 4 million) is located immediately to the north of eastern Australia. In 1987, it had a gross national product of per capita of $700(US), placing it 48th (equal with Swaziland) in the World Bank's ranking of poor countries.[7] Its population is predominantly Melanesian and it is governed from the coastal city of Port Moresby. Civil unrest and change in government have been common since independence. No national smoking prevalence surveys have been undertaken, although a 1981 survey conducted in two regions found rates as high as 85 percent among coastal Papuan men and 80 percent among highland women from Goroka.[8] The last published Household Expenditure Survey (1975/76) found that tobacco products accounted for 5.2 percent of average household expenditure, placing them after rent (10.5 percent), beer (7.2 percent), rice (5.6 percent), transport (5.3 percent) as the fifth most significant item in weekly expenditure.[9]

Smoking control policies and programs in Papua had been non-existent until the introduction of the Tobacco Products (Health Control) Act 1987 in November of that year.[10] The Act banned tobacco advertising; reportedly this ban has not been widely implemented. There are no health warnings on cigarette packs or advertisement, health education about smoking is negligible, and there has been neither public debate nor any regulation of smoking in public areas. Tobacco advertising and sports sponsorship both remain widespread, with the most remote bush trading posts typically being bedecked with posters advertising brands marketed by two Australian subsidiaries of transnational tobacco companies, i.e., W. D. and H. O. Wills and Rothmans. Cigarette advertising is almost certainly the most common form of outdoor advertising although no data are available on expenditure and consequently could not be included in our analysis.

Methods

Demand equations are usually estimated by regressing the quantity of the commodity demanded against the price of the commodity, the price of its close substitutes, and disposable income. Depending upon the commodity and the purpose of the analysis, other variables which influence demand may be included, such as advertising and health education expenditures.

In order to estimate the impact of taxes on consumption, a second equation must be estimated relating the level of taxation to the price level. This, in conjunction with the results of the demand equation may be used to infer the tax-consumption relationship. Regrettably, this conventional methodology could not be applied to the Papua New Guinea

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experience. Our inquiries about the contributions of wholesale and retail mark-ups to retail price were fruitless. Instead, the government statistical office in Port Moresby was able to provide information on the volume of cigarettes and non-cigarette tobacco consumed and the excise revenue obtained from these for the years 1973–86 (Table 1).

The use of excise rather than price data makes direct comparison of these findings with those obtained in other countries problematic although we believe our conclusions have important implications for policy since it is excise, and not usually retail price, that governments are able to manipulate directly. As excise tax is only a portion of retail price, excise elasticity is likely to be less than and so an *understatement* of the full price elasticity, thus suggesting that the elasticities shown below are quantitatively smaller than the relationship between retail price and demand. The two equations referred to above may be combined into a single equation which directly relates consumption to taxation. More formally, equations (1) and (2) below (which are usually estimated in demand studies) imply equation (3) which is the subject of the analysis here.

$$\text{Consumption} = f_1\,(\text{price},X) \qquad (1)$$

$$\text{Price} = f_2\,(\text{excise},Y) \qquad (2)$$

hence

$$\text{Consumption} = f_3(\text{excise},X,Y) \qquad (3)$$

where X and Y are sets of other relevant variables.
Equation (3) is a "reduced form" equation. Unlike the structural equations (1) and (2), it does not describe the direct, behavioral relationships between variables. More specifically, equation (1) embodies the fact that consumers react directly to market price and not to excise taxes. However, as retail price reflects taxes there is an indirect relationship which may be legitimately estimated. The resulting "excise elasticity" would not indicate the price elasticity of demand except under particular circumstances, i.e., when prices rise in direct proportion to taxes. If wholesale prices rose proportionately less than taxes (or remained unchanged or fell) then the percentage change in price would be less than the percentage change in tax and the excise elasticity would underestimate the magnitude of the price elasticity.

TABLE 1—Tobacco and Cigarette Consumption, Excise Tax Rate and Consumption Expenditure per Capita, Papua New Guinea 1973–86

Date	Quantity Tobacco kg/capita	Quantity cigarettes kg/capita	Excise Tax Rate Tobacco kina/kg	Excise Tax Rate Cigarettes kina/kg	Consumption Expenditure per Capita
1973	0.53	0.26	2.05	15.22	319.65
1974	0.39	0.22	2.45	20.03	263.56
1975	0.25	0.36	3.97	13.84	273.17
1976	0.25	0.33	4.85	16.51	281.66
1977	0.30	0.35	4.64	15.81	309.43
1978	0.29	0.38	4.46	15.31	326.11
1979	0.28	0.44	4.30	15.25	340.91
1980	0.26	0.44	4.05	13.45	355.07
1981	0.25	0.40	3.75	13.06	334.84
1982	0.27	0.34	3.53	15.78	313.69
1983	0.24	0.28	3.29	16.38	318.04
1984	0.26	0.30	3.07	16.04	317.48
1985	0.25	0.26	2.95	16.50	312.57
1986	0.27	0.26	2.80	16.37	—

The estimated "excise elasticity" would be biased if excise taxation correlated with one of the other variables and this variable was not included in the estimation procedure. The most important "other variables" in such analyses are the level of personal income per capita and the cost of production. As no information was available on personal income, data were obtained on a closely related measure, i.e., private consumption per capita. No data were available to act as a proxy measure for production costs. However, it appears most unlikely that these would correlate with annual changes in excise tax. Consequently, the regression coefficient should produce an unbiased estimate of the excise elasticity.

The data in Table 1 were used to estimate "demand" equations using ordinary least squares regression techniques and a multiplicative model shown below.

$$Q_I = e^{a+\beta 4T} E_I^{\beta 2} E_j^{\beta 2} C^{\beta 3}$$

where:
a = constant estimated by the regression
β1,β2,β3,β4 = coefficients estimated by the regression. Because of the multiplicative model, β1,β2,β3 are the elasticity coefficients
Q_I = quantity of product i per capita (in kg)
E_I = excise per kg of product i (inflation adjusted kina)
E_j = excise per kg of the substitute product j
C = total consumption expenditure per capita on all goods and services
T = a trend variable (1 in 1973; 2 in 1974; 2 in 1975 etc)
a_o = constant

The multiplicative model is used as it has simple but acceptable theoretical properties and it may be readily estimated with a logarithmic transformation of the variables.

The resulting coefficients β1-β3 may be interpreted as "excise elasticities" and "cross excise" elasticities. Thus, for example, in the equation for cigarettes, a 10 percent change in excise E_1 will be associated with a $10\beta_1$ percent change in the consumption of cigarettes (Q_1); a 10 percent change in the excise on the substitute, tobacco, will result in a $10\beta_2$ percent change in consumption. As it is the level of excise tax which may be directly varied by government and not generally consumer prices, the model is relevant for the policy issue of controlling consumption.

Results

Results for the two demand equations are reported in Table 2. A very high percentage of the variation in the quantity demanded is explained by the variables in the model as would be expected in a time series analysis. The "income" elasticity in both cases is positive and has a plausible magnitude. Surprisingly, the income effect for cigarettes is smaller than that for non-cigarette tobacco. As expected in the case of non-cigarette tobacco, there is a negative value for the trend variable. The coefficient suggests that, after allowing for the effect of the other variables in the model, non-cigarette tobacco consumption has declined about 5

TABLE 2—Regression Equation Results for the Quantity of Tobacco and Cigarettes per Capita

	Q (tobacco)		Q (cigarettes)	
Dependent Variable	B	SE	B	SE
Tobacco excise	−0.50*	0.08	0.50	0.09
Cigarette excise	0.62	0.22	−0.71	0.26
Income	1.37*	0.30	0.86	0.34
Trend	−0.05*	0.006	−0.008	0.006
Constant	−9.84	2.13	−4.68	2.45
R^2	0.95		0.93	
Standard Error	0.06		0.07	
dW	2.30		2.19	
n	14		14	

Notes: The first column of each result gives the estimated regression coefficient ( ) and the second column, the Standard Error. the logarithm of all variables has been used.

T (Trend variable) = 1, 2, 3, . . . 14

percent per annum.* No trend is apparent for cigarettes. The most important results in Table 2 are for the excise variables. For tobacco and cigarettes the excise elasticities are −0.50 and −0.71, respectively.** In both cases, excise "explains" a very high proportion of the variation in the quantity demanded. The results indicate that there was a strong relation between declining per capita cigarette consumption and rising excise. All else being equal, a 10 percent increase in cigarette excise has been associated with a 7.1 percent fall in cigarette consumption; and a 10 percent increase in non-cigarette tobacco excise with a 5 percent in its consumption. The "cross excise elasticities" (which due to the limited data have been unconstrained) are 0.62 for non-cigarette tobacco and 0.50 for cigarettes. That is, there is the expected positive sign in both cases.

Discussion

On the basis of the evidence presented here it would appear that the elasticity of demand for cigarettes and tobacco in Papua New Guinea is much greater than has been estimated for the USA and other Western nations in recent years. As excise is only a fraction of the final retail price, a change in excise would normally be expected to cause a proportionately smaller change in price which implies that the price elasticity exceeds the excise elasticity of demand. Irrespective of the relative magnitude of these coefficients, the present results suggest that the consumption of tobacco products can be manipulated through variations in the excise rate. If these demand characteristics are typical of developing countries, this finding may have enormous implications for smoking control in the Third World. Such countries often have diverse linguistic subcultures which mitigate against mass communication in health education. Being oriented toward problems of infectious disease and malnutrition, their health services typically accord low priority to smoking control. With the exception of a small number of developing countries in which tobacco is a major industry—e.g., Brazil, Zimbabwe, Malawi, Thailand, India—most developing nations have a deficit balance of trade in tobacco, with imports requiring the expatriation of growing amounts of hard currency to the exporting transnationals or state monopolies.[6]

Tobacco's main attraction to such economies lies in the ease with which its consumption can be taxed. However, excise taxes represent only transfer payments within economics whereas money lost overseas in paying for tobacco imports represents a genuine drain on a nation's balance of payments. As such, countries such as Papua New Guinea with deficit tobacco trade should not balk at easily implemented measures that reduce tobacco consumption if the resultant economic and public health advantages could be effectively communicated to key officials and "extraneous" factors like corruption and bribery exposed or otherwise circumvented.

In the short to medium term, an excise-led smoking control policy can reduce smoking and simultaneously satisfy Treasury requirements for what is an easily collected and bountiful source of revenue. Our results suggest that a 10 percent increase in cigarette excise will decrease consumption by about 7.1 percent, therefore increasing total excise collections. A 10 percent increase in tobacco excise will similarly increase total collections as the percentage increase in excise exceeds the percentage decrease in the volume consumed. Eventually, a threshold would be reached where the decline in consumption caused by rising prices would produce diminishing total excise. However, this point will, in many cases, be well in the future and therefore of less concern to political decision-makers than the attractions of both a short-term gain to state finances and falling smoking prevalence.

By World Bank projections,[7] the population of developing countries by year 2000 will constitute 80 percent of world population. Further, the aggregate rate of increase between 1987–2000 of the populations of developing countries is estimated to be 28 percent compared to a forecast increase of only 8 percent in the population of industrial market economies. Cigarette consumption is decreasing in the industrialized world at about 1 percent per year, whereas in Asia and Latin America it is growing 7 percent faster than the population and, in Africa, 18 percent faster. If the results of the excise elasticity analysis we have provided are found to be typical of those in developing countries, then excise policy could be a powerful instrument in controlling smoking in these vast markets. Its use could reduce tobacco consumption rapidly, possibly faster than any other realistic instrument of public health policy and hence produce health benefits for many millions of people.

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*The results are not the consequence of a correlation between income and trend. The coefficient of a correlation between these two variables is only 0.206. More generally, the correlation table does not suggest a problem of multicollinearity. (Data available on request from the authors.)

**This latter coefficient is increased to −1.08 if two further years (1971, 1972) are added to the data set. Excise data provided for these two years were outliers (33.98 and 34.49 kina/kg) and in the absence of any explanation for their large difference, we have omitted them from our main analysis, suspecting typographical error.

ACKNOWLEDGMENTS

We wish to acknowledge the very useful comments of an unknown reviewer and in particular his or her comments and calculations on the stability of the statistical results reported here.

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WHO Fellowships for Travel/Study Abroad

The World Health Organization (WHO) will provide a limited number of short-term fellowships for use during 1991 for travel/study abroad related to the improvement of health status in the United States. This support is limited to US citizens employed in the health field.

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The fellowship award provides per diem, transportation, and a small stipend to cover miscellaneous expenses. Employers of applicants are expected to endorse the application and agree to continue the applicant's salary throughout the fellowship period. Except in unusual circumstances, the fellowships will be limited to short-term programs of one or two months. The number of fellowship awards will be governed by the funds available.

The *deadline* for the submission of the completed application is September 30, 1990.

Further information and application form may be obtained from the Secretary, WHO Fellowships Selection Committee, Health Resources and Services Administration, 5600 Fishers Lane, Room 14-14, Rockville, MD 20857. Telephone: (301) 443-6152.

Smoking Before, During, and After Pregnancy

LOIS A. FINGERHUT, MA, JOEL C. KLEINMAN, PHD, AND JULIETTE S. KENDRICK, MD

Abstract: We report the first national data on smoking before, during, and after pregnancy. Estimates are based on the 1986 Linked Telephone Survey that reinterviewed 1,550 White women 20–44 years of age who were respondents to the 1985 National Health Interview Survey. An estimated 39 percent of White women who had smoked before pregnancy quit smoking while pregnant (27 percent when they found out they were pregnant and 12 percent later during pregnancy). Women with less than 12 years of education were five times as likely to smoke and one-fourth as likely to quit as those with 16 or more years of education. Women who smoked more than one pack of cigarettes per day before pregnancy were one-fifth as likely to quit as those smoking less. Of the women who quit, 70 percent resumed smoking within one year of delivery. Of those who relapsed, 67 percent resumed smoking within three months of delivery and 93 percent within six months. There is little evidence of educational differentials in relapse rates. The fact that relapse remains high suggests that while health of the fetus is a strong influence on women's smoking habits, women may be less aware of the effect of passive smoke on the infant. (*Am J Public Health* 1990; 80:541–544.)

Introduction

Increased risks of fetal and infant death and low birthweight have been linked to maternal smoking[1–3]; even if smoking is stopped late in pregnancy, there will be a positive effect on the infant's birth weight.[4,5] The effects of passive smoke on nonsmokers, and especially on young children, have also been shown to be harmful.[6–9]

For those women who smoke, pregnancy can be viewed as an ideal time for cessation: quitting can affect the health of their unborn child as well as their own health. It is also a time when contact with health professionals is very frequent. Thus there is substantial opportunity for health providers to have an impact on maternal and infant health through smoking cessation.

Among pregnant women, the proportion of married mothers 20 years of age and over who smoked before their pregnancy decreased between 1967 and 1980, while the proportion of those women who quit smoking after they found out they were pregnant increased.[10]

In this study, we report national data, the first ever available, regarding the frequency and timing of smoking relapse among women who quit smoking while pregnant.

Methods

Data Sources and Limitations

The National Health Interview Survey (NHIS) is a multistage stratified probability sample of the household population of the United States. In 1986, a linked telephone survey was conducted that reinterviewed respondents to the 1985 NHIS. The primary purpose of the telephone survey was to test methodology[11]; it included a battery of questions on smoking before, during, and after pregnancy. The target sample was selected from women responding to the NHIS during January–June 1985: nearly 2,000 women 18–44 years of age who either were pregnant at the time of the 1985 NHIS or whose most recent birth occurred during the five years preceding the NHIS. The women were interviewed using a computer-assisted telephone interview system during March–May 1986.

Definitions

- *Women who smoked prior to pregnancy* are those women who said that they smoked at least one cigarette per day during the 12 months before they found out they were pregnant. Women who said that they did not smoke regularly were not included as smokers prior to pregnancy.
- *Women who quit smoking while pregnant* are those who smoked prior to pregnancy and "quit early" (quit when they found out they were pregnant) or "quit later" (quit at a later time during their pregnancy). Those who quit early were not asked whether they resumed smoking prior to delivery. However, data from a large clinical trial of smoking during pregnancy show that very few women who quit early in pregnancy resume smoking before delivery.[12] Those who quit later continued to smoke cigarettes after they found out they were pregnant, but said that they quit for at least a week during their pregnancy and did not start smoking again before delivery; they were not questioned about when quitting occurred.
- *Women who relapse* are those who quit while pregnant and then resumed smoking within one year of the delivery.
- *Postpartum smokers* include women who relapsed within one year of delivery; women who continued smoking during the entire pregnancy and were still smoking within one year of the delivery; and women who smoked during the year after delivery even though they had not smoked in the 12 months prior to delivery.

The number of cigarettes smoked before pregnancy was dichotomized into less than 20 vs 20 or more cigarettes per day. Age of mother, education, and marital status were determined at the time of the 1985 NHIS.

Response Rates

The response rate to the first six months (January–June) of the 1985 NHIS was 96 percent.[13] Of these respondents, 86 percent completed the telephone followback survey yielding a sample of 1,918 women. Non-respondents included primarily refusals and those women with changed telephone numbers who could not be tracked.[11] In addition to excluding nonrespondents, 28 sample women who did not respond to

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the entire battery of smoking questions were omitted from the analysis.

As there were relatively few Black women in the sample (n = 278), this analysis is limited to White women. Further, there were only 35 White teenagers in the original sample and because teenagers are known to have different smoking habits than adults,[14] the analysis is limited to White women 20–44 years of age. The overall response rate for these women was 88 percent, yielding a final sample of 1,550 women.

For White women 20–44 years of age, response rates were higher for married than for unmarried women (90 vs 83 percent), and for those with at least a high school education than for those with less education (91 vs 78 percent). All percentages are weighted to reflect national estimates of women in the civilian noninstitutionalized population of the United States.

Analysis

The proportions and standard errors in this analysis were computed using the SESUDAAN program that takes into account the complex sample survey design of the NHIS.[15] Multiple logistic regression analysis was performed using the Proc Logist procedure within SAS.[16] Because the design effects were small (less than 1.2, based on the ratio of standard errors from SESUDAAN to simple random sample standard errors), standard errors were approximated using weighted estimates in the logistic regression procedure normalized to reflect the sample size (NORMWT option in SAS). The results of the logistic regression are presented as adjusted odds ratios with 95 percent confidence intervals calculated by taking the anti-logarithms of the regression coefficients and corresponding confidence intervals.

Results

One-third of White women 20–44 years of age said that they smoked cigarettes prior to pregnancy. Older women were less likely to smoke than younger women; and married women were much less likely than unmarried women to have been smokers. The prevalence of smoking decreased sharply with increasing education (Table 1).

Using multiple logistic regression to control for age and marital status, White women with less than 12 years education were five times as likely as college graduates (16 or more years education) to have smoked prior to pregnancy. Similarly, White unmarried women (controlling for age and education) were twice as likely as married women to have smoked. Women in their 20s were about 40 percent more likely to smoke than women in their 30s and early 40s (Table 1).

Of the White women who said they smoked, 39 percent indicated that they quit while pregnant (27 percent early and 12 percent later during pregnancy) (Table 2). The Health Promotion and Disease Prevention (HPDP) supplement to the 1985 NHIS estimated that 21.2 percent of all women 18–44 years quit smoking when they learned they were pregnant.[17] Further analysis of the HPDP data show that among White women 20–44 years of age, 24.2 percent quit smoking early in pregnancy. This estimate is not significantly different from our estimate of 27.4 percent.

Sixty percent of college graduates quit compared with 28 percent of those with less than 12 years education. Women who were lighter smokers prior to pregnancy were much more likely to quit than were heavier smokers.

Based on the logistic regression results, the relation between quitting and education was nearly as strong as between smoking and education. The amount smoked prior to pregnancy was also strongly associated with quitting. Age and marital status had little effect on quitting (Table 2).

In all, 70 percent of White women who quit smoking while pregnant resumed smoking within one year. This may be a slight underestimate since there were eight women who quit smoking while pregnant but who had less than a year of exposure time—from delivery to interview—in which to relapse and therefore were not considered relapsers. Unlike "smoking prior to pregnancy" and "quitting," there is little association between education and relapsing (Table 3). However, one year relapse rates are higher for those who quit later during pregnancy as opposed to those who quit when they found out they were pregnant (81 vs 66 percent).

Based on the regression results, younger women, women who were heavy smokers prior to pregnancy, and women who quit smoking later during pregnancy were more likely to relapse. However, the relatively small numbers of women who quit smoking (n = 191) result in rather wide

TABLE 1—Prevalence of Smoking Prior to Pregnancy among White Women 20–44 Years of Age

	Sample Size*	Percent Smoking	Standard Error	Adjusted Odds Ratios** (95% CI)
All women	1550	32.0	1.3	—
Age (years)				
20–29	832	37.5	1.7	1.38 (1.09, 1.74)
30–44	718	25.1	1.7	ref
Marital Status				
Unmarried	127	52.5	5.1	2.01 (1.37, 2.94)
Married	1423	30.1	1.5	ref
Education (years)				
Less than 12	165	54.1	4.3	4.99 (3.26, 7.65)
12	679	37.4	1.9	2.71 (1.94, 3.77)
13–15	353	26.7	2.4	1.73 (1.19, 2.52)
16 or more	352	16.1	1.9	ref

*Sample size is number of White women 20–44 years of age. Education was missing for one woman.

**Based on multiple logistic regression.

TABLE 2—Smoking Cessation during Pregnancy among White Women 20–44 Years of Age

	Sample Size*	Percent Quitting	Standard Error	Adjusted Odds Ratios** (95% CI)
All women	486	39.3	2.3	—
Age (years)				
20–29	308	38.3	2.8	1.05 (0.68, 1.61)
30–44	178	41.1	4.0	ref
Marital Status				
Unmarried	64	36.6	7.2	0.88 (0.49, 1.60)
Married	422	39.7	2.4	ref
Education (years)				
Less than 12	86	27.5	4.9	0.26 (0.12, 0.56)
12	249	37.0	3.0	0.39 (0.21, 0.75)
13–15	93	44.4	5.5	0.51 (0.25, 1.07)
16 or more	57	60.1	7.0	ref
Amount smoked before pregnancy				
<1 pack/day	204	59.0	3.6	ref
1+ packs/day	281	24.6	2.7	0.22 (0.15, 0.33)

*Sample size is number of women who smoked prior to pregnancy. Education and amount smoked were missing for one woman.

**Based on multiple logistic regression.

TABLE 3—One-year Relapse among White Women 20–44 Years of Age Who Quit Smoking while Pregnant

	Sample Size*	Percent Relapsing	Standard Error	Adjusted Odds Ratios** (95% CI)
All women	191	70.5	3.6	—
Age (years)				
20–29	121	74.0	4.0	1.94 (0.95, 3.94)
30–44	70	64.4	7.0	ref
Marital Status				
Unmarried	24	64.3	9.1	0.72 (0.27, 1.90)
Married	167	71.4	3.8	ref
Education (years)				
Less than 12	25	58.2	10.1	0.36 (0.10, 1.22)
12	92	75.2	5.7	1.05 (0.42, 2.66)
13–15	41	68.9	7.1	0.73 (0.25, 2.10)
16 or more	33	68.5	7.9	ref
Amount smoked before pregnancy				
<1 pack/day	121	66.6	4.5	ref
1+ packs/day	69	76.8	5.3	1.58 (0.77, 3.25)
Quit Time				
Early	131	65.9	4.6	ref
Later during pregnancy	60	81.0	5.4	2.15 (0.96, 4.82)

*Sample size is number of women who quit smoking while pregnant. Amount smoked was missing for one woman.

**Based on multiple logistic regression.

confidence intervals for all of these estimates (Table 3). Age of mother and time of quitting (early vs later) were the only factors for which the confidence intervals came close to excluding 1.0.

Of all White women who quit smoking while pregnant, 27 percent relapsed within a month of the delivery, and 52 percent relapsed within 4 months (Figure 1). Those women who quit later during pregnancy resumed smoking much sooner than women who quit smoking earlier in their pregnancy; 69 percent of the former group relapsed within three months of the delivery compared with 37 percent of the latter group of women.

The net effect of quitting and relapsing is the prevalence of smoking one year postpartum. The estimate includes women who relapsed within a year, those who did not quit during their pregnancy, and those who smoked postpartum but said they were not smoking before or during their pregnancy. (This latter group is small—only 2 percent of the total sample.) The great majority, 87 percent, of the women who smoked prior to pregnancy were also smokers one year after delivery. Hence, the prevalence of smoking one year after the delivery was only slightly lower than before pregnancy; 30.3 vs 32.0 percent. Postpartum smoking has similar associations with age, marital status and education as shown in Table 1 for smoking prior to pregnancy.

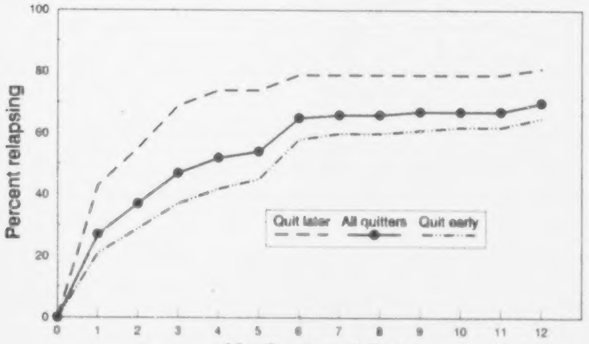

FIGURE 1—Relapse Rates among White Women Ages 20–44 Years Who Quit Smoking While Pregnant (data are from the 1986 Linked Telephone Survey)

Prenatal Care and Smoking

Nearly all (99 percent) of the women said that they received prenatal care during their pregnancy. Those who smoked prior to pregnancy were less likely to begin prenatal care in the first trimester than those who did not smoke during this period (90 percent vs 95 percent). Those smokers who quit while pregnant were more likely to have begun care in the first trimester than those who did not quit, 94 vs 88 percent.

Regardless of whether a woman went to a clinic or to a private physician for prenatal care, about three-fourths of the smokers prior to pregnancy said that a doctor had advised them to quit or cut down on smoking during their pregnancy (57 percent who quit smoking and 83 percent who did not quit).

In response to a question about different methods used to quit or help cut down on smoking, about 95 percent of the women who quit smoking stated that they quit entirely on their own, without a formal program or counseling. While some of the women also said that they stopped with the help of friends and that they switched to low tar and nicotine cigarettes to help themselves quit, clearly the vast majority, regardless of whether they quit early or later during pregnancy, attributed their behavior to their own initiatives.

Discussion

The potential effect of recall bias on the results of this study should be considered. Because of the survey design, a woman may have been asked to recall her smoking habits that occurred five or six years earlier. About 22 percent of the women in the sample delivered in 1980–81, 56 percent from 1982–84, and 22 percent in 1985–86. There was a decrease in the proportion of White women smokers between 1980 and 1985–86, similar to the decline observed in the NHIS data collected during those years.[18] Moreover, smoking history recall over a 10-year period was found to be quite reliable.[19] Thus, recall bias in the current study should not have a major effect on the conclusions.

Nearly all women in this sample received prenatal care and most smokers reported that they had been advised by a doctor to quit smoking. Those who did not quit were, in fact, more likely than quitters to report such advice (83 vs 57 percent); this apparent paradox probably reflects their longer opportunity to be given this advice on repeated prenatal visits. We cannot distinguish the role that counseling during prenatal care might have played in encouraging women to quit; it is possible that women who quit early did so in response to counseling that occurred at their first prenatal visit when they were told they were pregnant. Several previous studies have documented that the advice of a doctor or nurse can be very important in encouraging smokers to quit.[20] A variety of materials are available to health providers who want to provide smoking cessation help to their smoking patients.[21]

We found strong educational differentials in smoking prevalence and in quitting. This reinforces the conclusions of a recent study of smoking prevalence in the general population.[22] Between 1974 and 1985 smoking declined five times faster among college graduates compared with those with less than a high school education.

Pregnant women appear to quit smoking at a greater rate than the general population. Based on data from the 1986 Adult Use of Tobacco Survey, an estimated 28 percent of persons 17 years of age and over who had smoked in the preceding 12 months quit smoking for at least one week during the year.[1] The proportion of White women who quit for at least a week while pregnant is higher—39 percent.

Relapse rates among women who quit smoking during pregnancy appear to be somewhat less than among quitters in the total population. In general, about 80 percent of smokers who quit resume smoking within a year. This is similar to the relapse rates for other drug dependencies.[23] In our study, 81 percent of White women who quit late in pregnancy also resumed within a year, but only 66 percent of those who quit early resumed smoking again within a year. Thus for some women, pregnancy appears to be an event that could be more likely to result in longer-term abstinence than smoking cessation for other reasons.

It is also noteworthy that women who quit smoking while pregnant resume smoking after delivery at a similar pace to quitters in the general population. In both instances, there is a precipitous rise in the proportion who resume smoking by three months followed by a relatively smaller increase by six months. The likelihood of resuming after six months is less than during the preceding months.[23] The underlying difference between these two groups, however, is that these pregnant women were abstinent during their pregnancy before they resumed smoking again.

However, the fact that relapse rates are high suggests that while the fetus's health is a strong influence on women's smoking habits, women may be less aware of the effect of passive smoke on the infant. Children of parents who smoked are more likely to have a variety of health problems such as respiratory illnesses than children of parents who do not smoke.[6–9] Education about the deleterious effects of infant and childhood exposure to passive smoke should be an essential part of comprehensive prenatal and postpartum care. Intervention programs for pregnant smokers should focus on assisting them to quit permanently.

ACKNOWLEDGMENTS

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Screening for *Chlamydia Trachomatis* in Adolescent Males: A Cost-Based Decision Analysis

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Abstract: To evaluate the cost and benefits of screening tests for *Chlamydia trachomatis* in adolescent males, we developed a decision analysis model and compared the leukocyte esterase urine dipstick test with culture, with direct-smear fluorescent antibody (DFA), and with the option of no screening (no treatment). The leukocyte esterase test has the lowest average cost-per-cure ($51) compared with direct-smear fluorescent antibody ($192) and culture ($414). Compared with the DFA, we estimate that the leukocyte esterase test saves over $9,727 per cohort of 1,000 sexually active adolescent males screened. Sensitivity analyses show the leukocyte esterase test results in a lower cost-per-cure and lower overall costs (per cohort) than culture and direct-smear fluorescent antibody at any prevalence of *C. trachomatis* infection, and lower overall costs (per cohort) than no screening at prevalences above 21 percent. (*Am J Public Health* 1990; 80:545–550.)

Introduction

Infections caused by *Chlamydia trachomatis* are among the most prevalent and costly in the United States. Over four million chlamydial infections occur each year at a cost exceeding $1.5 billion dollars.[1,2] Adolescents have the highest rates of chlamydial infection and associated complications of any age group.[3–5] Considerable attention is now being focused on young women at risk for contracting *C. trachomatis* because of the dire reproductive consequences of a chlamydial infection. These include mucopurulent cervicitis, acute pelvic inflammatory disease, ectopic pregnancy, infertility, and maternal and infant infections during pregnancy and following delivery.[6–8]

Young men are equally at risk for a chlamydial infection. *C. trachomatis* causes approximately 50 percent of the reported cases of nongonococcal urethritis among males, and is also responsible for over 50 percent of cases of acute epididymitis in young men[8]; epididymitis is a painful, serious complication of chlamydial infection that could result in sterility. Moreover, infected adolescent males, particularly the 30 percent with asymptomatic chlamydial urethritis, represent the major source of *C. trachomatis* transmission to teenage females.[4,9]

Nevertheless, limited resources and limited availability of diagnostic tests have dictated the need to focus on screening females first. With newer, inexpensive and reliable tests now available, however, routine screening of males should be considered and evaluated. In this article, we evaluate the cost and benefits of the three commercially available screening methods for *C. trachomatis* in a population of adolescent males at risk for infection. We compare the leukocyte esterase urine dipstick test to urethral culture, to antigen detection* of a urethral smear, and with the option of not screening or treating.

Methods

Decision Analysis Model

We used a decision tree (Figure 1) to evaluate the possible outcomes associated with three screening tests for *C. trachomatis* in a population of adolescent males: 1) leukocyte esterase (LE) urine dipstick with a 1+ or 2+ reading on first catch urine; 2) urethral culture; and 3) antigen detection* (DFA) or a urethral smear. These were compared with the option of no screening (no treatment). Baseline probabilities used in our model were obtained from published data (Table 1). Our denominator for calculating baseline outcomes and costs is a cohort of 1,000 sexually active adolescent males.

Data and Assumptions

Chlamydia Prevalence (Node 1): We used 15 percent as the baseline prevalence to be conservative in estimating outcomes, but we also performed sensitivity analyses** to test the effect of using different prevalences on our conclusions.[4,9–14]

Sensitivity (Node 2) and Specificity (Node 3) of Screening Test: We used a conservative sensitivity estimate of 75 percent.[10,11] Since many false positive dipstick tests will be due to the presence of *N. gonorrhoeae*, for which at least 85 percent of strains are susceptible to tetracycline,[15] and has sequelae for uncured cases similar to that of *C. trachomatis*, we assumed a specificity of 85 percent based on results of screening only asymptomatic males.[11] Test sensitivity and test specificity for culture of *C. trachomatis* were assumed to be 90 percent and 99 percent, respectively.[16] For the DFA test we estimated the sensitivity to be 80 percent and the specificity to be 97 percent.[17,18] We performed a sensitivity analysis of the LE test through the range of probabilities in Table 1.

Follow Up (Node 4): When there is a delay between taking a test and the availability of results, some patients will be lost-to-follow-up. We used a loss-to-follow-up rate of 3 percent (range 3 to 10 percent) for the culture and direct-smear (DFA) test, because in most clinical settings results are not available for one to three days after obtaining a specimen. This follow-up rate is based on a high yield, cost-effective, field follow-up method of tracking males with positive chlamydial tests.[19] Results from the LE urine dipstick are

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*Direct-smear fluorescent antibody (DFA) examination. Commercially available as Microtrak (tradename), manufactured by Syva Diagnostics.

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**Sensitivity analysis is the process of varying assumptions in a decision analysis.

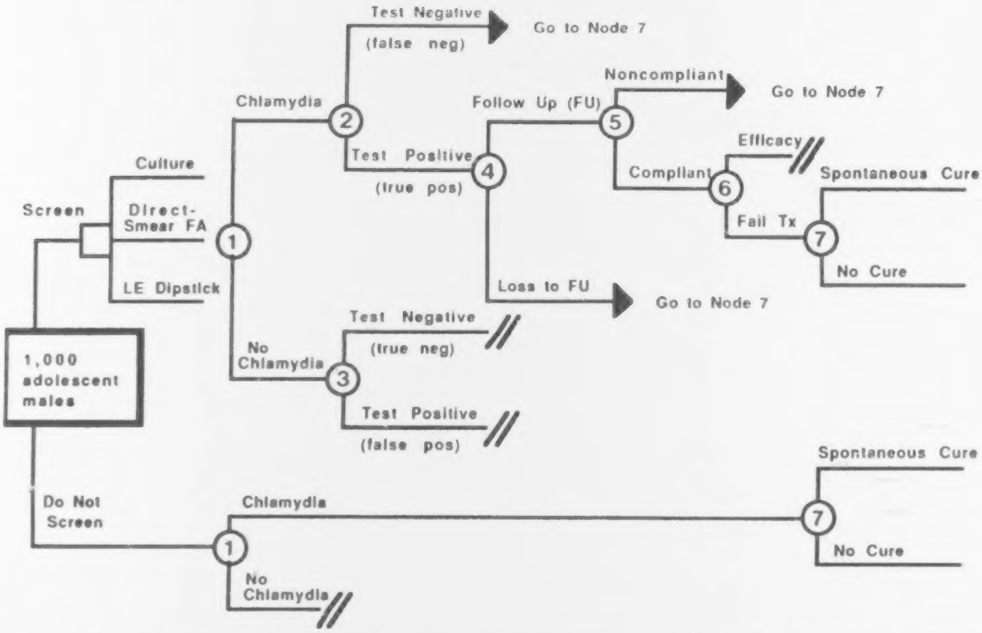
Test Negative
(false neg)
Go to Node 7
Noncompliant
Go to Node 7
Chlamydia
Follow Up (FU)
Efficacy
Test Positive
(true pos)
Compliant
Spontaneous Cure
Culture
Screen
Direct-Smear FA
Fail Tx
No Cure
LE Dipstick
Loss to FU
Go to Node 7
Test Negative
(true neg)
1,000 adolescent males
No Chlamydia
Test Positive
(false pos)
Spontaneous Cure
Do Not Screen
Chlamydia
No Cure
No Chlamydia
1
2
3
4
5
6
7
1
7

FIGURE 1—Decision Tree Depicting Outcomes for Screening (with three different tests) or not Screening for *Chlamydia trachomatis* in Adolescent Males. Squares indicate decision nodes; circles, chance nodes; and asterisk, sexually active adolescent males.

TABLE 1—Probabilities Used in Decision Analysis

	Probability		
	Baseline %	Range %	Reference
Prevalence of chlamydia	15	5–35	4,9–14
Sensitivity of leukocyte esterase	75	65–65	10,11
Specificity of leukocyte esterase	85	75–95	10,11
Sensitivity of culture	90		16
Specificity of culture	99		16
Sensitivity of direct-smear fluorescent antibody	80		17,18
Specificity of direct-smear fluorescent antibody	97		17,18
Loss-to-follow-up rate	3	3–10	19
Compliance with treatment	65	65–95	20–22
Effectiveness of treatment	95		15,23
Spontaneous cure rate	5		8
Pelvic inflammatory disease rate (infected partners)	20	10–30	3,26,29
Probability of infecting a female	30	20–40	19

generally available immediately, therefore we assumed no loss-to-follow-up for this screening test.

Compliance (Node 5): We conservatively estimated the compliance to be 65 percent (range 65 to 95 percent) in this population based on studies that show: a compliance rate of 66 to 79 percent of men treated with a five-day course of doxycycline for *N. gonorrhoeae*,[20] and an overall compliance rate for pediatric populations of approximately 50 percent, with a range of 20 to 80 percent.[21,22] We assumed that noncompliers had failed treatment, and that only those with spontaneous cure (5 percent) were free of disease.

Efficacy of Treatment (Node 6): Cure rates for the tetracyclines in treatment of chlamydial urethritis in males range from 91 percent to 99 percent.[15,23] We used a cure rate of 95 percent for our base-line analysis, the mean rate from published reports. Theoretically, other drug regimens such as trimethaprim-sulfamethoxasole[15] could also be used with the cost estimates adjusted accordingly.

Spontaneous Cure Rate (Node 7): We assumed a spontaneous cure rate of 5 percent for infected males who did not receive treatment.[8]

Cost Estimates

This decision model used medical care costs as utilities (Table 2). Only direct costs were considered with charges used to approximate costs.

Screening Test: We surveyed several laboratories in the San Francisco Bay metropolitan area to derive an average cost of $0.50 for LE dipstick (including the cost of the urine container), $30 for culture, and $10 for DFA. The cost of collecting the tests, performing the LE urine dipstick in the office, and of packaging the DFA and culture and sending them to the laboratory was assessed as $10.

Follow-up and Loss-to-Follow-up: The cost of tracking down clients with positive cultures or positive DFA tests and arranging an office visit has been shown to be $10.67.[19] Since a fair amount of time can be spent attempting to track down

TABLE 2—Cost Assumptions Used in Decision Analysis

Leukocyte esterase urine dipstick	$ 0.50
Culture	30.00
Direct-smear fluorescent antibody	10.00
Test collection and processing	10.00
Follow-up	10.67
Loss-to-follow-up	10.67
Treatment (doxycycline hyclate)	10.00
Complications in untreated males	34.34
Infecting a female partner	366.72

persons lost to follow-up, a cost of $10.67 was included for all clients with positive cultures or DFAs lost to follow-up.

Treatment: The $10 cost of treatment includes the cost of medication and the pharmacy charge. The current bulk cost for a seven-day treatment course of doxycycline hyclate 100 mg PO bid is less than $2.[24] We added in an additional $8 charge for dispensing, storing, and ordering the medication.

Complications in Sex Partners: The average cost of infecting a female partner, including transmission to neonates, was $366.72 using the baseline PID rate of 20 percent (Appendix Table A). This estimate is based on the probabilities shown in Table 1, and costs per incident of medical care in Table 2, and published data on the probability of chlamydial complications in women and neonates and associated costs.[1,2,25,26] The Appendix provides details of these calculations. This cost was decreased to $195.38 per female by decreasing the pelvic inflammatory disease (PID) rate to 10 percent, and increased to $538.05 per female by increasing the PID rate to 30 percent. We assumed that the probability of a male infecting a female was 0.30.[19] However, this variable was also ranged from 0.20 to 0.40.

Complications in Males: The cost of return visits for symptomatic infections in untreated males was assessed to be $45 ($35 office visit plus $10 treatment) with a probability of occurrence of 75 percent. This means that only about 56 percent (1×.75×.75) of infected men will return, because 25 percent will be asymptomatic. The cost of outpatient treatment for epididymitis was assessed to be $50 with a probability of occurrence of 3.6 percent in untreated males.[2,8] The cost of inpatient treatment for epididymitis was calculated to be $1,836 with a probability of occurrence in 0.4 percent of untreated males.[2,8] This results in a total cost for complications per untreated male of $34.34.

Results

Cost Savings at Baseline Assumptions

When total medical costs and cost-per-cure are considered, screening with the LE test results in the greatest cost savings (Table 3). Screening with the LE test costs 12 percent as much as culture and 27 percent as much as the DFA. The average cost-per-cure of a chlamydial infection is $51 when the LE test is used compared with $414 per cure with culture and $192 per cure with the DFA. This cost-per-cure is unchanged by decreasing the cohort of adolescent males screened to 100. Overall, screening with the LE test will save over $28,459 per 1,000 adolescent males at risk for *C. trachomatis* infection compared with culture and $9,727 compared with the DFA.

TABLE 3—Overall Cost and Outcomes for Different Screening Strategies*

	Cost		Health Outcome	
Screening Strategy	Total	Per Cure**	Cure (%)	Unnecessary Rx
Leukocyte esterase	$23,944	$ 51	73 (49%)	219
Culture	52,403	414	84 (56%)	15
Direct-smear fluorescent antibody	33,671	192	76 (51%)	44
No screening	20,571	—	8 (5%)	0

*Per cohort of 1,000 sexually active adolescent males screened for *C. trachomatis* infection.

**Cost-per-cure = (total costs screening method minus total costs no screening)/(total cures screening strategy minus total cures no screening).

TABLE 4—Sensitivity Analysis on Seven Variables

		Cost/Cure	
Variables	Percent	LE	DFA
Prevalence			
	35	$ (−)53	$ 19
	5	$ 415	$ 795
LE Sensitivity/Specificity			
	85/95	$ 19	$ 192
	65/75	$ 93	$ 192
PID rate in infected partners			
	30	$ 0	$ 140
	10	$ 103	$ 243
Probability of infecting a female			
	40	$ 14	$ 155
	20	$ 88	$ 229
Compliance			
	95	$ (−)11	$ 86
Loss-to-Folloow-up DFA			
	10	$ 51	$ 216
Best Case*		$(−)195	$(−)137
Worst Case**		$ 619	$ 866

* = Prevalance = 35%; LE Sensitivity 85%/Specificity 95%; PID Rate 30%; Infected Partner Rate 40%; Compliance 95%

** = Prevalence = 5%; LE Sensitivity 65%/Specificity 75%; PID Rate 10%; Infected Partner Rate 20%; Compliance 65%.

Health Outcomes at Baseline Assumptions

Screening with culture results in the highest percent cure rate (56 percent) per 1,000 adolescent males (Table 3). Screening with the DFA results in the second highest percent cure rate (51 percent), while the LE results in the lowest percent cure rate of the three tests (49 percent). The strategy of not screening results in a 44 to 51 percent lower cure rate than any of the three screening options. Fewer patients will be treated unnecessarily with culture and direct-smear (DFA) screening compared with the LE test (see Table 3).

Sensitivity Analysis and Break-Even Prevalence

Sensitivity analyses were performed on seven variables (Table 4). Because the LE test and the DFA appear to be the more practical options for screening, culture was not included in the sensitivity analysis.

At base-line assumptions, the disease prevalence which leads the LE test to break even when compared with the cost of not screening is 21 percent. Above this prevalence, the LE test always costs less than not screening. The break even prevalence for the DFA is 41 percent. Using base-line assumptions, the LE test will be less costly than the DFA at any prevalence. As the prevalence is lowered to 5 percent, the cost/cure for the LE rises to $415 and for the DFA to $795.

Increasing the sensitivity/specificity of the LE test will, as expected, make the LE test even more cost-effective (lower cost-per-cure, and/or lower total costs and higher or equivalent cure rate) compared with the DFA. Even at the lowest sensitivity/specificity for the LE, the cost-per-cure is still only 48 percent of the cost of the DFA.

Increasing the PID rate in infected partners to 30 percent results in the LE breaking even compared to not screening. The DFA would still cost $140 per cure. Decreasing the PID rate to 10 percent results in the LE costing $103 per cure with the DFA costing $243 per cure. Increasing the probability of infecting a female to 40 percent results in the LE test costing $14 per cure and the DFA costing $155 per cure. Decreasing the probability of infecting a female partner to 20 percent results in the LE costing $88 per cure and the DFA costing $229 per cure.

If the loss-to-follow-up for culture and the DFA is raised to 10 percent, the percent cure rate for culture is still the highest (53 percent), with the LE next (49 percent) and the DFA the least effective (47 percent cure). The cost per cure for the DFA rises to $216 compared to $51 for the LE. If the loss-to-follow-up rate is 7 percent, the percent cure rate for the LE is equal to the DFA.

If the compliance rate were increased to 95 percent with all other variables at base-line, the LE would save $11 per cure (cost-per-cure = $51 at 65 percent compliance) and the cost-per-cure for the DFA would drop to $86 (from $192 at 65 percent compliance).

If the "best case" scenario for use of screening tests is applied (high sensitivity/specificity of tests, high prevalence of disease, high PID rate in infected partners, high probability of infecting a partner) the LE saves $195 per cure, the DFA saves $137 per cure, and culture saves $6 per cure. If the "worst case" scenario for use of screening tests is applied (low sensitivity/specificity of tests, low prevalence of disease, low PID rate in infected partners, low probability of infecting a partner) then the LE costs $619 per cure and the DFA costs $866 per cure.

Discussion

Our analysis indicates that the leukocyte esterase (LE) urine dipstick test, when used to screen for *Chlamydia trachomatis* in a population of sexually active adolescent males results in greater cost savings than screening with either DFA or culture. This finding is true throughout the range of sensitivity analyses for seven separate variables. Moreover, screening with the LE urine dipstick test in this population can result in an equivalent number of cures if the loss-to-follow-up rate for the DFA test is greater than 6 percent. These findings clearly underscore the effectiveness and substantial benefits of the LE test and support its use in this at risk population that is currently not routinely screened.

In populations where contacting patients with test results and treatment recommendations is not problematic and the prevalence of *C. trachomatis* is very high, the cost-per-cure for the DFA is substantially reduced. The higher costs associated with use of the DFA over the LE must be balanced against the high incidence of unnecessary treatment resulting from the lower specificity of the LE test. Since the LE false positive rate was established using cultures which probably have <90 percent sensitivity,[17] the LE may be detecting infections which would be missed by DFA or culture. We also recognize that some false positive LE dipstick tests in adolescent males are due to the presence of *N. gonorrhoeae* so that treatment unnecessary for the eradication of chlamydia would ultimately be of benefit. We did not attempt to place a value (or utility) on the outcome of unnecessary treatment. While the morbidity associated with tetracycline therapy in males is certainly minimal, the psychosocial consequences of inappropriate diagnosis and unnecessary treatment could conceivably impose a heavy burden on these young males and their sex partners.

Recognizing that the specificity of the LE test is much lower than that estimated for the DFA, and that precise identification of an infecting organism has potential therapeutic and health education value, have led practitioners to advocate the use of the LE as a screening test leading to further testing with culture or DFA.[10,11] Although use of this strategy would decrease the number of unnecessary treatments, the number of cured chlamydial infections would be lowered due to the combined false negative rates of both tests. Overall costs and cost-per-cure would be increased when compared with use of the LE alone. In populations where the prevalence of chlamydial infection is high and few patients are lost to follow-up, the cost-per-cure for the DFA may be acceptable enough for use as a screening test. In these populations, the higher expense for the DFA compared with the LE may be offset by the decreased number of unnecessary treated patients. Use of the LE as a predictor for further testing may be most beneficial in populations with a low prevalence of chlamydial infection and a high rate of follow-up, since the low specificity of the LE has the greatest relative impact in terms of unnecessary treatments at low disease prevalences.

Compliance with treatment is an important variable that significantly impacts the total costs and cost-per-cure for each screening strategy. If compliance with treatment could be increased to 95 percent, screening with the LE would save money over no screening. Similarly, increasing the compliance with treatment significantly reduces the cost-per-cure for the DFA and, if the follow-up rate was also high, some practitioners may find the cost-per-cure for the DFA to be in an acceptable range even at moderately low prevalences.

Our analysis focused on screening only for *C. trachomatis* with the LE test when the studies we cite[10,11] which determined the effectiveness of the LE test actually addressed both *C. trachomatis* and *N. gonorrhoeae* in asymptomatic adolescent males. The sensitivities and 100 percent specificities found in these studies were based on confirmed infection with either *C. trachomatis* or *N. gonorrhoeae*, or both. Since the sensitivity was 100 percent in one study, and the specificity was 100 percent in the other, our conservative sensitivity and specificity assumptions of 75 percent and 85 percent for the LE should absorb the effect of *N. gonorrhoeae*.

As with the LE, the sensitivity and specificity of the DFA in this population of all sexually active adolescent males are unknown. The DFA test was evaluated in men with symptoms or signs of urethritis and has not been evaluated in low-risk, asymptomatic men.[17] The sensitivity and specificity may be lower in adolescent male asymptomatic carriers. The LE test is aimed at detecting the pool of asymptomatic adolescent males who may be the main transmitters of infection to adolescent females. Sensitivity and specificity of the DFA test are also highly dependent on the skill and expertise of the microscopist interpreting the slide.[17] In laboratories with inexperienced technicians, the number of false-positive and false-negative results will bring the sensitivity and specificity toward the lower end of the spectrum.

We did not include in our analysis the strategy of empiric treatment of all sexually active adolescent males attending a clinic or physician's office. One study of patients at a sexually transmitted disease (STD) clinic[27] found this strategy to be the most cost-effective when compared with the strategies of culturing all patients and only treating those with positive results; empirically treating patients at high risk and culturing those at low risk; and empirically treating patients at high risk and neither culturing nor treating those at low risk. We initially evaluated this strategy and found it resulted in cost-per-cure of $(-)14, with 95 cures, 55 no cures, and 850 unnecessary treatments per 1,000 adolescent males. However, we elected not to present the empiric treatment data in our comparative decision analyses because we agree with other investigators that this is not a feasible health policy option.[27] Not only would there be no basis for determining frequency of treatment, except perhaps level of sexual activity, but deciding about treatment of sex partners would also prove difficult. Furthermore, indiscriminate use of tetracycline could adversely affect antimicrobial sensitivity patterns, leading to resistant *N. gonorrhoeae* and possibly *C. trachomatis* as well.

The major costs associated with adolescent male chlamydial infection result from the costs associated with infecting a female partner. Although we have data to support the incidence of sequelae from chlamydial infection in females, and the associated costs resulting from the medical management of these sequelae, there are no firm data on the transmission rate from males to their female sexual partners. Recognizing that the transmission might have been from female partner to the adolescent male, that some adolescent females might already be infected with the same serovar of *C. trachomatis* from another partner and cannot be reinfected, and that the transmission rate is probably lower than 100 percent, we conservatively assumed that an infected male has a 0.30 change of infecting a female. Studies have shown that 30 percent of reported female partners of males with urethritis will be infected with *C. trachomatis*.[19] Since many adolescents have multiple sexual partners and may not report all of these partners, our sensitivity analysis which ranges the probability of infecting a female from 0.20 to 0.40 should cover any inaccuracy surrounding this variable.

These considerations notwithstanding, our analysis provides reasonable estimates of the risks, benefits, and costs of screening for *C. trachomatis* with the LE urine dipstick test compared with culture, DFA and not screening. Where imprecision exists, we believe it is in the direction of underestimating benefits and cost savings. For example, we did not include indirect costs, which are about equal to direct costs[2,26]; adding them would further increase the cost saving and enhance the cost-effectiveness of the LE test. Neither did we include the noneconomic benefits. These include the prevention of adverse psychosocial effects in males at risk for epididymitis and in their female partners who might later suffer from PID, ectopic pregnancy (with its resultant fetal loss and risk of maternal mortality), or infertility.

One other benefit not included in our analysis is the treatment of coexisting *N. gonorrhoeae* in adolescents diagnosed with *C. trachomatis* by the LE test. While data are not available on this coinfection rate, the rate of coexisting *C. trachomatis* in patients diagnosed with *N. gonorrhoeae* ranges from 15 to 30 percent in males and 25 to 50 percent in females.[8,13,28] Assuming the rate of coexisting *N. gonorrhoeae* in patients diagnosed with *C. trachomatis* is similar, then a significant number of gonococcal infections will be treated along with the chlamydial infection, because a seven-day course of doxycycline hyclate (100 mg PO bid) will eradicate at least 85 percent of infections due to *N. gonorrhoeae* since the highest reported rate of tetracycline resistent *N. gonorrhoeae* is 15 percent.[15,20] Importantly, treatment of these asymptomatic gonococcal infections would prevent transmission to susceptible females and subsequent adverse reproductive consequences. This would lead to even greater cost savings with the LE test compared to other screening methods and to not screening than what was calculated in our study.

The added benefits of the LE in detecting *N. gonorrhoeae* leads to combined *C. trachomatis/N. gonorrhoeae* treatment recommendations.[29,30] Not only does there appear to be an increasing treatment failure rate among patients with *N. gonorrhoeae* taking tetracycline exclusively,[31] but also the expected compliance with a seven-day treatment course is only 65 percent. Compliance with *N. gonorrhoeae* treatment would be raised to 100 percent if LE dipstick-positive adolescent males were given 250 mg ceftriaxone intramuscularly once or a one time oral dose of 3.5 g ampicillin/1 g probenicid before leaving the physician's office, and then given their seven-day course of doxycycline hyclate 100 mg PO bid to eradicate 95 percent of *C. trachomatis* in those who are compliant with treatment. Tracking of penicillin-resistant *N. gonorrhoeae* could be accomplished by culturing all males with a discharge, dysuria, history of exposure to a sexually transmitted infection, and/or a positive LE urine dipstick test for *N. gonorrhoeae*.

Although there are several studies that compare the costs and benefits of screening programs for *C. trachomatis* in females,[27,32,33] this is the first study that directly compares the costs and benefits of screening in adolescent males. This population represents a key link in the chain of transmission of chlamydial disease, yet routine screening of this group is not widely practiced. Since the cost-per-cure is unaffected by the size of the cohort, our analysis provides a model that practitioners can apply once they outline the characteristics of their cohort. To prevent further transmission and minimize the risk that an infected individual will have adverse consequences, clinicians should assess the prevalence and loss-to-follow-up rate in their populations, and consider using the LE urine dipstick test or the DFA, alone or in combination, for screening sexually active adolescent males to detect *Chlamydia trachomatis* infection.

APPENDIX

Cost-effectiveness Calculations

The cost of infecting a single female partner and the costs of potential neonatal transmission and its probability are calculated as shown in Appendix Table A. Appendix Table B shows the cost per *C. trachomatis* infected male assuming that 30 percent (range 20 to 40 percent) of female partners exposed by the male would become infected and have the costs and risks in Table A. We did not include potential male-to-male transmission. We assumed that 20 percent of female partners would present for treatment.[19] The cost of treating infected female partners included two outpatient visits, one DFA test, a seven-day course of doxycycline and treatment of the vaginitis expected in 10 percent of treated partners from the antibiotic therapy. Given the 20 percent false negative rate for the DFA,[17,18] we assumed that 16 percent of infected female partners would be treated with complete eradication of *C. trachomatis* and prevention of any sequelae. For the 84 percent of untreated female partners, we expected a 20 percent PID rate (range 10 to 30 percent) with the costs of treating the PID and it's sequelae listed in Table A. We predicted a 5 percent birth rate and estimated a 20 percent chlamydial conjunctivitis rate and a 10 percent chlamydial pneumonia rate in neonates born to *C. trachomatis* infected mothers. The cost of complications in males infected with chlamydia is combined with the cost of infecting female partners and the costs of potential neonatal transmission to result in an average cost per *C. trachomatis* infected male of $144.36 (Table B). The number of infected males was calculated by multiplying probability estimates down each pathway of the model and considering all those ending with "no cure" as infected and all those ending with efficacious treatment or spontaneous cures as "cured." Total costs were calculated by multiplying all "no cure" totals with the cost of infected males ($144.36) and adding this product to the following: 1) the cost of each test times the cohort; 2) the cost of follow-up and loss-to-follow-up times their probabilities of occurrence; plus 3) the cost of treatment and its probability of occurrence. This total cost calculation was divided by the number "cured" to obtain the cost-per-cure. The number unnecessarily treated equals the number of false positive tests. We used a microcomputer spreadsheet program to model the number of cures and costs.

APPENDIX TABLE A—Cost of Infecting Female Partners with *C. trachomatis* and Transmission to Neonates

	Cost	Probability	Reference
Infected female partners			
Rx fluorescent antibody		.20	
Direct-smear	$ 8	1.00	
Two outpatient visits	70	1.00	25
Vaginitis (complication Rx)	30	.10	26,27,29
7-Day course doxycycline	10	.80	
Infected female partners not Rx		.84	
PID Rate		.20	3,26,29
PID-outpatient treatment	180	1.00	25,26,29
PID-inpatient treatment	4,259	.21	25,26,29
Surgery	1,500	.12	25,26,29
Ectopic pregnancy	5,759	.04	25,26,29
Tubal infertility	2,500	.21	25,26,29
Chronic pelvic pain	200	.15	25,26,29
Birth Rate		.05	
Neonatal pneumonia	1,375	.10	7,8,25,29
Neonatal conjunctivitis	55	.20	7,8,25,29
Average cost/case	365.12		

APPENDIX TABLE B—Cost per *C. trachomatis* Infected Male

	Cost	Probability	Reference
Infected female partner & neonatal transmission	$ 365.12	0.3	Table A
Return visit symptoms	45.00	0.56	27
Epididymitis outpatient treatment	50.00	0.036	2,8
Epididymitis inpatient treatment	1,836.00	0.004	2,8
Average Cost/Case	144.36		

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Information on Tar and Nicotine Yields on Cigarette Packages

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Abstract: We examined information on tar and nicotine yields on the packages of 160 cigarette brands, 58 percent of the 275 brands for which tar and nicotine yields were listed in a recent Federal Trade Commission report. The tar yield was indicated on 14 percent, the nicotine yield was indicated on 11 percent. As tar yield increased among brands, the yield was progressively less likely to be shown on the package and was not disclosed on the package of any cigarette yielding 11 mg or more of tar. (*Am J Public Health* 1990; 80:551—553.)

Introduction

Since 1971, all major cigarette manufacturers have voluntarily disclosed the tar and nicotine yields of cigarette brands in advertisements.[1] The cigarette industry agreed to "voluntary" disclosure after the US Federal Trade Commission (FTC) had proposed a regulation that would have required such disclosure.[2] This agreement does not apply to cigarette packages.

The tar and nicotine yields listed in cigarette ads have been obtained by machine smoking of cigarettes under standardized laboratory conditions using an FTC method developed in the 1960s.[3] Through early 1987, the FTC conducted tar and nicotine measurements in its in-house laboratory; currently, the Tobacco Institute Testing Laboratory performs tar and nicotine measurements using the FTC's standard method. The Institute provides these test results to the FTC which, in turn, makes the information available in a public report.[4]

In 1981, the Secretary of Health and Human Services recommended that:

> ". . . manufacturers should list yields of 'tar', nicotine and other hazardous components on their packages and in their advertising with appropriate explanatory information on the health significance of these measurements. This would be a minimum first step in giving cigarette consumers full and adequate information about the products they are buying."[5]

Seven years after that recommendation was made, we conducted a study to examine the types of information related to tar and nicotine yields that are provided on cigarette packages.

Methods

In May of 1988, we visited eight retail stores that sell cigarettes in Montgomery County, Maryland to find as many different cigarette brands as possible. We found 160 (58 percent) of the 275 brands listed in a 1988 FTC report on tar, nicotine, and carbon monoxide yields of domestic cigarettes.[4] (The 275 brands include 272 in the report's primary listing, plus three varieties of Barclay cigarettes for which tar and nicotine yields were provided in the text.) These 160 brands accounted for approximately 78 percent of the domestic cigarette market in the second quarter of 1988.[6]

Many of the 160 brands were "extensions" of the same primary brand name but differed in cigarette length, the type of package (soft or hard pack), the number of cigarettes per package (20 or 25), the presence or absence of a filter, or the presence or absence of metholation. All of these variables (including type of package) may be associated with differences in tar and nicotine yield.[4,7]

For each brand, we examined the package for any information relating to tar and nicotine yields, including quantitative measurements (mg per cigarette) and qualitative descriptions (e.g., "light," "low-tar," "99% tar free"). Because tar and nicotine yields in cigarette brands are highly correlated (high-tar brands tend to have high nicotine yield, and low-tar brands tend to have low nicotine yield),[4] we analyzed package labeling patterns only with respect to tar yields.

For six brands, the tar yield indicated on the package was 1–2 mg higher or lower than the yield listed in the FTC report. In these cases, we used the yield indicated on the package in our analysis of package labeling patterns.

Results

Package Disclosure of Tar and Nicotine Yields

The tar yield was indicated on the package of 23 brands (14 percent). For 17 (74 percent) of the 23 brands, the nicotine yield was also indicated. In no case was the nicotine yield shown without the tar yield. The 23 brands fell within six brand "families" produced by five different manufacturers: seven Carlton varieties (American Brands); seven Merit varieties (Philip Morris Company); three Now varieties (RJR/Nabisco); three Barclay varieties (Brown & Williamson Corp.); two Triumph varieties (Loews Corp.); and one Benson & Hedges variety (Philip Morris). The 23 brands accounted for approximately 5 percent of the domestic cigarette market in the second quarter of 1988,[6] or about 6 percent of the combined market share of the 160 brands in our sample (78 percent).

Table 1 shows the proportion of brands whose packages disclosed tar yield within the six tar-yield categories used by the FTC in analyzing cigarette sales and advertising expenditures.[8] As tar yield increased among brands, the package was progressively less likely to show the yield. The tar yield was disclosed for all 11 brands yielding 3 mg or less of tar but not for any of the brands yielding 11 mg or more. When the six tar-yield categories shown in Table 1 are collapsed into four categories to allow use of the chi-square test for trend, the trend is statistically significant ($p < 0.001$, two-tailed).

Package Label Messages

Table 2 shows the frequency of use of words and phrases on cigarette packages that indicate or suggest a low-tar/

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TABLE 1—Proportion of Cigarette Brands That Disclose Tar Yield on Packages, among a Sample of 160 Brands (stratified by tar yield), United States, May 1988

Tar Yield* (mg/cigarette)	Total Brands	Brands with Tar Yield on Package	
		N	%
≤ 3	11	11	100
4–6	19	6	32
7–9	29	5	17
10–12	46	1	2
13–15	23	0	0
≥ 16	32	0	0
Total	160	23	14

*Source of information on tar yields of cigarette brands is the Federal Trade Commission.[4]

low-nicotine yield, stratified by tar yield. Five brands claimed to have the lowest yield. Two varieties of Carlton cigarettes, each having a tar yield of 1 mg, claimed that "Carlton is lowest." Three varieties of Now cigarettes, having tar yields of 1 mg, 2 mg, and 3 mg, claimed that "Now is lowest." Three varieties of Barclay cigarettes, ranging from 3 to 5 mg of tar, were said to be "99% tar free."

The phrases "ultra low tar" and "ultra lights" were used almost exclusively for brands yielding 7 mg or less of tar. The terms "low tar" and "lights" were generally used for brands yielding 15 mg or less of tar, conforming with the FTC's usual definition of low-tar cigarettes.[9] However, three brands with tar yields of 16–18 mg were called "lights" (Doral Lights king-size soft-pack, Richland Lights 100's soft-pack, and Richland Lights Menthol king-size soft-pack). One "ultra low tar" brand (Merit Menthol 100's soft pack) had a tar yield of 10 mg, which is higher than that of some brands labeled as merely "low tar." One brand (Lucky Strike Unfiltered Lights) had the phrase "only 9 mg tar" on the package.

Several words or phrases that may suggest a low-tar yield are used on packages, including "smooth," "mild," "mellow," and references to special filters (Table 2). One filter was described in detail on the package: "Lark Lights filter has two outer 'tar' and nicotine filters, plus an inner chamber of tiny charcoal granules specially treated to smooth the taste in a way no other filter can." The term "full flavor," often used by cigarette companies as a euphemism for high tar, appeared on the package of three brands yielding 14–17 mg of tar (Cambridge and Doral varieties).

TABLE 2—Frequency of Use of Words and Phrases on Cigarette Packages That Indicate or Suggest a Low-Tar/Low-Nicotine Yield, among a Sample of 160 Brands (stratified by tar yield), United States, May 1988

Tar Yield* (mg/cigarette)	Wording	Brands	
		N**	%
≤ 6 (N = 30)	Ultra low tar	16	53
	Ultra lights	13	43
	Lowest	5	17
	99% tar free	3	10
	Ultra low	2	7
	Smooth	2	7
7–12 (N = 75)	Light/Lights	49	65
	Low tar	29	39
	Lowered tar & nicotine	5	7
	Ultra low tar	3	4
	Mild/Milds	3	4
	Low nicotine	2	3
	Smooth	2	3
	Ultra lights	1	1
	Recessed filter	1	1
	Famous Micronite II filter	1	2
≥ 13 (N = 55)	Light/Lights	8	15
	Smooth	6	11
	Low tar	5	9
	Mellow	2	4
	Activated charcoal filter	2	4
	Famous Micronite II filter	1	2
	Mild	1	2

*Source of information on tar yields of cigarette brands is the Federal Trade Commission.[4]

**Within a particular tar-yield category, the sum may exceed the total because some package labels used more than one descriptive term.

Discussion

Although several studies have shown that those who smoke lower tar cigarettes have a reduced risk of lung cancer,[5,10,11] many concerns have been expressed about low-yield cigarettes.[12,13] First, because the cigarettes are less harsh, they may facilitate experimentation with and initiation of smoking among children and adolescents. Second, smokers who switch to lower yield brands often change their smoking behavior to compensate for reduced nicotine delivery (e.g., by increasing daily cigarette consumption; by increasing the depth, duration, or frequency of puffs; or by smoking cigarettes to a shorter butt length).[14–17] These compensatory changes may, in some smokers, increase the total intake of tobacco smoke constituents. Third, the use of flavoring additives is thought to be more common in low-tar brands to compensate for reduced "taste." These additives and their products of combustion may pose additional health risks to smokers. Fourth, those who smoke lower yield cigarettes may consider them to be safe or less hazardous, and may therefore be less motivated to quit.[18] Fifth, there is little or no evidence that low-yield cigarettes reduce the risk of cancers other than lung, of cardiovascular disease, of chronic obstructive pulmonary disease, or of fetal damage when smoked during pregnancy.[5,13,19]

Consumers have received mixed messages about the advisability of switching to low-tar cigarettes. For example, widespread publicity in the media has followed the publication of articles reporting "benefits" of low-yield cigarettes.[20] These reports have been buttressed by cigarette advertising claims implying that low-tar cigarettes are less hazardous or safe.[18] On the other hand, articles reporting no "benefits" have also received substantial publicity.[19]

Consumers' ability to sort out and act on these disparate messages is complicated by the fact that package information on tar and nicotine is provided selectively and is ambiguous or inconsistent at times. Among the 160 brands included in our study, the tar yield was not disclosed for any brand yielding 11 mg or more of tar. (The sales-weighted average tar yield was 13 mg in 1987.[21]) Five different brands claimed to be "lowest," even though the tar yield of one (3 mg) was three times the tar yield of another (1 mg). One "ultra low tar" brand had a tar yield higher than that of several "low tar" brands. Three brands with tar yields ranging from 3 to 5 mg claimed to be "99% tar free," an ambiguous term which suggests the brands are virtually free of tar.

Requiring disclosure of tar and nicotine yields on all cigarette packages would eliminate the selective manner by which cigarette companies now provide this information. Uniform disclosure on packages would also make these figures more accessible than those currently provided on cigarette advertisements. A standard coding system should be developed to define tar-yield categories using easily understood terms (e.g., low tar, medium tar, high tar). The appropriate designation could be required on each package and advertisement. Such a coding system is used in the United Kingdom for cigarette advertisements.[22] In the United States, no official government coding system exists for tar-yield terminology, and no industry standard apparently exists (Judith P. Wilkenfeld, Federal Trade Commission, oral communication, July 6, 1989).

Chapman, *et al*,[22] have suggested that retailers be required to display at the point of sale a table showing the yields of all brands (from lowest to highest yields). According to these authors, the state of South Australia has adopted a law implementing this recommendation.

Above all, mechanisms should be developed to warn consumers that any health benefits of switching to lower tar brands are minimal when compared with the benefits of quitting and are offset by smoking more cigarettes per day, puffing more frequently, etc. Warnings could be required, for instance, on package inserts, in advertisements, and at the point of sale. According to the 1986 Adult Use of Tobacco Survey, 21 percent of smokers (approximately 10 million Americans) believe that the kind of cigarettes they smoke are less hazardous than other cigarettes.[21] Many of these smokers might quit if they understood that any benefits of switching to low-yield cigarettes pale in comparison to the benefits of quitting.

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Comparative Evaluation of American Cancer Society and American Lung Association Smoking Cessation Clinics

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Abstract: We compared the effectiveness of American Cancer Society, *FreshStart*, American Lung Association, *Freedom from Smoking*, and laboratory clinic methods in subjects (N = 1041) from three communities. Three-month follow-up results favored the laboratory method over the two public service approaches on both a prevalence and a sustained abstinence measure. At one-year follow-up, treatment effects for smoking prevalence were no longer significant. However, sustained abstinence results at one-year remained highly significant and favored the more intensive laboratory and *Freedom from Smoking* clinics over the *FreshStart* method. *FreshStart* fared less well than the other interventions both in producing initial quit attempts and in sustaining abstinence among initial quitters. It should be noted, however, that *FreshStart* requires considerably less facilitator contact than do the other approaches. Unexpected outcome effects occurred for treatment location. Future clinic programs should include a specific target date for quitting and should place more emphasis upon recycling participants who fail to sustain abstinence. (*Am J Public Health* 1990; 80:554–559.)

Introduction

Public service programs have been developed to assist chronic smokers to quit. Characteristics common to most of these programs include not only the use of behavioral principles, but also a focus on multiple program components, increasing emphasis on maintenance and relapse prevention, and a heightened appreciation of variables such as cognitive processes and social support.[1] As smoking cessation programs have become relatively more sophisticated, there has also been an increased awareness of the need for more cost-effective delivery methods.[2] Among the most widely offered public service programs are smoking cessation clinics developed by the American Cancer Society and the American Lung Association. Despite their widespread use, no evaluations of the effectiveness of these clinics have appeared to date in the literature.

The present study undertook a comparative evaluation of the American Cancer Society (ACS) *FreshStart* and American Lung Association (ALA) *Freedom from Smoking* clinic methods. These methods were assessed in a context that also included a laboratory-derived clinic approach.[3,4] The laboratory method was included to allow not only assessment of the effectiveness of the two public service clinics themselves, but of their effectiveness relative to that of a method systematically derived from a paradigmatic program of research[5] which has been systematically disseminated in community settings as part of a program of technology transfer. Costs have been kept to a minimum through the use of donated facilities and volunteer group leaders who receive only nominal fees.[6]

Methods

Subjects

Subjects were 1,041 smokers (57.2 percent women) who responded to newspaper advertisements announcing the availability of free smoking cessation clinics. The clinics were described as being part of a federally funded project. Subjects were recruited from three Iowa locations: Des Moines (N = 743), Iowa City (N = 173), and Waterloo-Cedar Falls (N = 125). Three study locations were used to ensure adequate subject enrollment and to allow a test of the major study hypotheses. The investigator in charge of administering treatment in Waterloo-Cedar Falls left after the first year of the study and recruitment efforts in the area were discontinued at that point.

Facilitators

Facilitators were 10 women and five men specifically recruited and trained to administer intervention for purposes of the current study. Although specific information on facilitator characteristics is not available, group leaders in Des Moines were older, ranging in age between 35 and 50, and most had previous experience in conducting group programs. In contrast, facilitators in Waterloo and Iowa City consisted primarily of undergraduate and graduate students less than 30 years old. Efforts were made to see that each facilitator conducted all three interventions. This was not always possible, however. It should be emphasized that we have no evidence of a facilitator effect upon treatment outcome.

All facilitators were formally trained in all three methods: the American Cancer Society *FreshStart* program, the American Lung Association *Freedom from Smoking* clinics, and the Lando laboratory derived approach. *FreshStart* training consisted of a four-hour program provided by the American Cancer Society. The smoking intervention coordinator for the American Lung Association of Iowa provided training in the *Freedom from Smoking* method. Project staff provided training in the Lando method. Formal instruction in the American Lung Association and Lando clinic procedures required approximately six to eight hours. Facilitators were provided copies of detailed treatment manuals for all three clinic approaches.

Facilitators were informed that the study was intended to assess the relative effectiveness of different methods, that it was unclear whether any method would prove superior to the others, and that a major purpose of the study was to determine which methods might be more effective for different people. Emphasis was placed upon the point that a critical overriding study goal was to assist as many people as possible to quit across all conditions. Every effort was made to minimize bias on the part of group leaders. Facilitators were recruited from each of the communities and individuals were employed who had no prior identification with any of

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the programs. Strong emphasis was placed upon the need to conduct procedures exactly as described in the respective protocols.

Experimental Design

Subjects were randomly assigned to one of the three methods as a function of orientation sessions attended. Subjects were also informed that the study was comparing different methods and that while each of the methods was effective, the investigators believed that a particular method might be more effective for certain types of smokers. Treatment was offered free of charge.

Assigning all individuals attending a given orientation session to a single method minimized problems of self-selection to treatment. This was deemed important enough to outweigh any limitations arising from having different units of randomization and analysis. Only very general information that could apply to any of the three methods (e.g., small group format and a varying number of sessions) was released prior to the orientation. Friends and relatives who came together to orientation sessions were ensured the same treatment condition. A total of 70 orientation sessions were held and 97 treatment groups were eventually formed in the three intervention sites.

Preintervention Assessment

Subjects completed a brief demographic and smoking history questionnaire as well as an informed consent statement at the initial orientation session and provided the names and telephone numbers of three informants who could assist in locating them for purposes of follow-up data collection.

All subjects were informed that the programs would teach them skills to actively confront urges to smoke and enable them not only to stop smoking, but to remain nonsmokers. They were asked to verbalize a commitment to quitting that was intrinsic rather than a product of coercion from others. Emphasis was placed upon the premise that smoking is a learned habit and that it can be unlearned as well.

Attendance records allowed an indication of attrition *before* the fact (e.g., differential dropout subsequent to the orientation sessions as a function of anticipated treatment procedures). Details of the different interventions are contained in the Appendix.

Follow-up Data

A brief smoking status questionnaire was mailed to subjects at three-month, six-month, and 12-month follow-up intervals. Subjects who failed to respond to mailed questionnaires were contacted by telephone. Point prevalence was defined as complete abstinence from tobacco products on the day in question. Sustained abstinence was defined as the absence of relapse. Permitted slips were distinguished from relapses following guidelines recommended by the National Heart, Lung, and Blood Institute National Working Conference on Smoking Relapse[7] (a relapse was defined as seven consecutive days of smoking at least one puff per day). We specified an additional definition of relapse as seven consecutive slips within three months. Data were collected on whether subjects had made a quit attempt during the clinic they attended. A quit attempt was defined as at least 24 hours of continuous tobacco abstinence.[7]

Eighty subjects who claimed sustained abstinence through one-year follow-up (43.0 percent of the total claiming sustained abstinence) were asked to submit saliva samples about one month after their one-year follow-up had been completed. Saliva samples were collected in subjects' homes. All samples were analyzed for thiocyanate. Samples with borderline values (80–100 ng/ml) or higher (which appeared to contradict self-reported abstinence) were subsequently analyzed for cotinine (smoking was defined as a value of greater than 15 ng/ml). In the case of subjects who refused to submit to saliva testing, informants were telephoned and queried concerning the subjects' smoking status.

Results

Subject Characteristics

The average age of subjects was 42.4 years. Mean baseline cigarette consumption was 28.8 per day and the average number of years smoking was 23.6. Subjects had initiated a mean of 3.8 prior quit attempts with the longest abstinence period averaging 7.7 months. In an open-ended question asking about reasons for attempting to quit, 86 percent of subjects listed health concerns as one reason for quitting (Table 1).

Multivariate analysis of variance, performed to identify location differences, indicated that the Iowa City sample was on average 6.9 years younger and smoked 4.4 cigarettes less than their counterparts in the other two communities. Clinic size (i.e., the number of participants per clinic) was also considerably smaller in Iowa City.

Response Rates at Follow-Up

There was a 93.8 percent response rate at one-year follow-up. Response rates at three-month and six-month follow-ups were 98.2 percent and 93.0 percent, respectively. Forty-two people were not successfully contacted, 15 were lost to follow-up, five refused the follow-up interview, and three died. Those not reached, lost to follow-up, and refusers

TABLE 1—Mean Values on Baseline Variables for Total Sample and by Each Location Separately

Measure	Total	Des Moines	Iowa City	Waterloo
Sample Size	1041	743	173	125
Age	42.4 (±12.7)	43.7 (±13.1)	36.3 (±10.4)	43.1 (±10.4)
% Females	57.3	58.5	57.8	48.8
Cigarette consumption	28.8 (±12.2)	29.4 (±12.4)	25.3 (±10.9)	30.4 (±12.3)
Cigarette nicotine level	0.75 (±.30)	0.76 (±.31)	0.72 (±.26)	0.72 (±.30)
Years smoking	23.6 (±12.0)	24.7 (±12.5)	18.2 (±9.6)	25.2 (±10.4)
Previous quit attempts*	2	2	2	3
Longest quit attempt (months)*	1.1	1.1	1.3	1.2
% giving health reason to quit	86.1	86.7	81.7	88.4
Self-efficacy (0–100)	53.4 (±19.2)	52.4 (±19.1)	54.5 (±19.3)	57.4 (±19.2)
Clinic Size	10.7 (±4.0)	11.8 (±3.6)	7.2 (±2.2)	12.5 (±5.0)

*The median is reported for this variable because of the substantial skewness of responses.

were counted as smokers for purposes of data analysis. There were no differences between the treatment groups or locations in attrition rates at one-year.

Return to Treatment

For 62 of 70 informational meetings, data were available on the number attending the meeting as well as the number who returned to participate in treatment. A weighted least squares ANOVA indicated that the proportion returning was not a significant function of the treatment being offered.

Group size did not differ significantly as a function of the intervention method. Mean group sizes were 10.0, 11.7, and 10.5 for the ACS, ALA, and Lando methods, respectively.

Abstinence

The effects of method and location (and their interaction) on abstinence outcome were evaluated using logit analysis.[8] Lack of independence of outcome within treatment groups was expected due to individuals receiving treatment together in groups and having correlated outcomes, and related individuals (e.g., spouses) entering the same treatment group. A group effect was observed on one-year abstinence prevalence after method and location effects were taken into account ($\chi^2(92) = 123.32$, $p = .016$). Thus, an estimate of the design effect[9] due to clustering was computed (= 1.469), and the original chi-square test statistics were divided by the design effect to correct for this within group dependence.[10] Corrected chi-square statistics and confidence intervals are reported below.

Location effects did not interact with those due to treatment classification in any of the analyses reported. Also there were neither facilitator nor sex of subject effects on abstinence outcome.

Abstinence outcomes are illustrated in Figure 1; mean levels are presented in Table 2. While point prevalence differences are no longer statistically significant after the three-month follow-up interval, sustained abstinence differences are significant at all follow-up points. The ACS *FreshStart* program fared most poorly, and the Lando program achieved better outcome than the ALA *Freedom from Smoking* method.

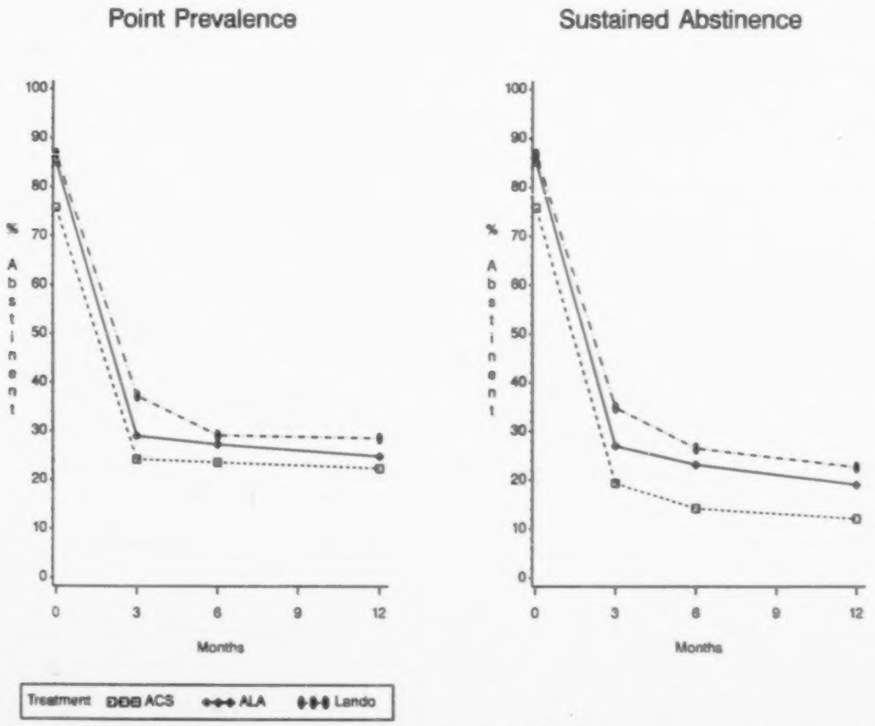
Ninety-five percent confidence intervals comparing the ACS *FreshStart* treatment with the average of the two more intensive programs, and the Lando with the ALA *Freedom from Smoking* program on various follow-up measures are presented in Table 3. The *FreshStart* program fared less well than the other two methods on all measures.

Point prevalence abstinence rates at three months were highest in Waterloo (41.6 percent), followed by Des Moines (29.7 percent) and Iowa City (23.7 percent). These differences no longer held at six months and one year. However, at one-year follow-up, sustained abstinence rates were 27.2 percent, 17.0 percent, and 15.0 percent in Waterloo, Des Moines, and Iowa City, respectively ($\chi^2(2) = 6.05$, $p = .048$).

Quit Attempts

There were significant differences in the proportion of subjects in the different treatment groups making quit attempts (Table 4). Subjects in the ACS *FreshStart* program were again less likely to make a quit attempt than those in the other two programs.

Considering only those who made a quit attempt, there was no significant difference in one-year point prevalence


Point Prevalence

Sustained Abstinence

FIGURE 1—Abstinence Outcome at Three Follow-up Times Using Two Abstinence Criteria (Initial abstinence rates (at months = 0) represent initial quitting rates, i.e., those who remained abstinent for 24 hours after quitting.)

TABLE 2—Outcome Rates (expressed in percent) at Three Follow-up Intervals

Outcome Measure	ACS	ALA	Lando	Total	P*
3M Point Prevalence	24.17	28.93	37.18	30.16	.005
6M Point Prevalence	23.56	27.27	29.11	26.71	.38
1 Yr Point Prevalence	22.36	24.79	28.53	25.26	.28
3M Sustained Abstinence	19.34	27.00	34.87	27.19	<.001
6M Sustained Abstinence	14.20	23.14	26.51	21.42	.004
1 Yr Sustained Abstinence	12.08	19.01	22.19	17.87	.014
Total N	331	363	347	1041	

*P value is derived from χ^2 statistic with 2 degrees of freedom corrected for DEFF = 1.469.

TABLE 3—Percentage Differences and 95% Confidence Intervals for Percentage Differences in Outcome Measures among the Treatment Conditions

	ACS vs Others		Lando vs ALA	
Outcome Measure	Difference*	95% CI	Difference	95% CI
3M Point Prevalence	−8.76	(−16.02, −1.51)	10.37	(1.61, 19.14)
6M Point Prevalence	−4.45	(−11.41, 2.51)	2.73	(−5.51, 10.98)
1 Yr Point Prevalence	−4.20	(−11.00, 2.60)	4.33	(−3.76, 12.42)
3M Sustained Abstinence	−11.68	(−18.56, −4.80)	9.70	(1.07, 18.33)
6M Sustained Abstinence	−10.90	(−17.25, −4.56)	4.82	(−3.36, 13.00)
1 Yr Sustained Abstinence	−8.85	(−14.92, −2.79)	4.94	(−2.90, 12.78)

*Estimates of differences and confidence intervals are adjusted for location effects.

TABLE 4—Quit Attempts and One-Year Outcome among Those Making a Quit Attempt

Outcome Measure	ACS	ALA	Lando	Total	P*
Making quit attempt	75.83	85.12	86.74	82.90	.004
1 Yr Point Prevalence	27.09	28.48	31.56	29.15	.59
1 Yr Sustained Abstinence	15.94	22.33	25.58	21.60	.068

*P value is derived from χ^2 statistic with 2 degrees of freedom corrected for DEFF = 1.469.

outcome between the methods but persons in the ACS *FreshStart* program were less likely to sustain abstinence than were those in the other two treatment groups [difference = −8.02%; 95% confidence interval = (−15.37, −0.67)].

Clinic location did not affect the number making quit attempts but did influence one-year sustained abstinence outcome among those attempting cessation (data available on request to author).

Biochemical Validation

Results of biochemical validations performed at one-year tended to be consistent with self-reported abstinence. Of the 80 persons selected for biochemical validation, 73 provided saliva samples. No samples exceeded the cutoff for disconfirmation. Six of the remaining seven persons, while claiming abstinence, refused biochemical validation while one person had returned to smoking between the one-year follow-up and testing. One of those who refused was identified as a smoker by an informant and three additional refusers either did not list or we were unable to contact their informants. If all four of these subjects are counted as smokers (which appears to be an appropriate, albeit conservative, practice) the lie rate would be 5.1 percent (= 4/79). Adjusting the reported outcome data by this factor would not markedly affect absolute abstinence levels and would have little impact upon interpretation of results.

Discussion

The present study represented the first comparative evaluation of the American Cancer Society *FreshStart* and the American Lung Association *Freedom from Smoking* clinic methods. Results indicated relatively modest outcomes for both clinics. The disappointing outcome for the *FreshStart* method on the sustained abstinence measure (12.1 percent sustained abstinence at 12 months) appears noteworthy. This figure would be even lower (10.9 percent) if more rigorous continuous abstinence guidelines had been used (i.e., absolutely no use of tobacco in any form whatsoever).

The reduced number of initial quit attempts with the *FreshStart* method would appear to be largely a function of the absence of a set quit date. However, the present study design does not permit unequivocally eliminating other possible explanations including the lesser total clinic contact. Even so, it should be noted that intensity of clinic contact prior to quit attempts was fairly comparable for the *FreshStart* and *Freedom from Smoking* methods. The major difference in scheduling between the two public service approaches occurred during the maintenance phase rather than in preparation for quitting. Although the results for *FreshStart* are disappointing, it should be remembered that the *FreshStart* method also requires considerably less facilitator contact.

The fact that these programs were offered totally free to participants may have been partially responsible for overall modest abstinence rates. The absence of a fee may also have reduced participants' commitment.

Facilitator bias cannot be dismissed as a possible factor mediating differences between methods. However, every effort had been made to eliminate this bias. Facilitators had no a priori commitment to a given approach. They had been specifically instructed to work within the confines of each method to produce the best possible outcomes. Furthermore, the larger differences occurred between the two public service methods. Differences between *Freedom from Smoking* and the laboratory program (in which the investigators would have been expected to have a vested interest) were less pronounced and were no longer significant at the one-year follow-up.

Differential enrollment in treatment conditions was not a factor in mediating outcome. There were no differences in enrollment in the three clinic programs following the initial informational meeting. It is possible that a limitation inherent in the interpretation of these results may arise from the fact that the unit of randomization and the unit of analysis in this study are different. However, the large sample size provides a measure of confidence in the results.

The location effects were unanticipated. Although no obvious explanation is apparent for the superior outcome in Waterloo, clinics were relatively new to this community. Numerous cessation programs had been offered in both Des Moines and Iowa City. Perhaps smokers in Waterloo responded to the novelty of the clinic programs. Des Moines and Iowa City subjects would have had many previous opportunities to enroll in cessation clinics and these subjects may have been more resistant to quitting. It should be noted that at no point did location effects interact with intervention. The order of outcomes was comparable across all three study communities despite the main effects for location. The fact that the study results were replicated in three separate communities with a relatively large total subject pool lends increased confidence in their generalizability. Ironically, the Iowa City sample which fared most poorly was younger and apparently less habituated (as indicated by lower daily cigarette consumption) than were the other subject samples.

Although cost-effectiveness analyses were not a formal part of the present study, cost considerations are obviously important.[11–12] Costs of programs can be sharply reduced by use of lay volunteers and donated facilities.[6] Training requirements are minimal. Each of the three methods can be learned in a few hours of formal instruction. Costs that are incurred can be largely defrayed by relatively modest fees to participants (the maximum fee per participant in public service applications of the laboratory-derived clinic method is $75.00). Costs should be lower for the ACS method given the requirement of only four hours of participant contact. These lower costs must be weighed against the reduced effectiveness of this method.

It has proven possible to achieve low-cost technology transfer of the laboratory method. Collaboration between laboratory researchers and a voluntary agency (the American Lung Association of Iowa) has proven critical in effecting this transfer. Especially encouraging is the apparent lack of decay in outcome in the transfer between the laboratory and field settings. There have been virtually no net costs to field sites (often community hospitals). Participant fees have defrayed costs of facilitators and promotion (often promotion has occurred through public service messages).

The American Lung Association of Iowa in the 1988–89 fiscal year reported total expenses of $32,438.70 in overseeing the statewide implementation of 111 laboratory-derived public service clinics reaching approximately 1,000 smokers. Expenses included research and evaluation, travel, facilitator training and updates, printing, postage, computer and software, telephone, personnel, advertising and audiovisuals. These expenses were largely offset by income ($18,235.00), fees charged for facilitator training and updates ($2,982.08), and a grant from an insurance company ($3,187.34). Net expenses were $8,034.28 or approximately $8.03 for each smoker who went through the program. Cost per abstinent smoker at 12 months was $26.77. Considerable public health impact might be achieved through widespread application of this type of collaborative model involving voluntary agency dissemination of validated cessation clinic programs.

The disappointing overall success rates suggest the need for further intervention among non-abstinent subjects. Prevalence outcomes at one year considerably exceeded sustained abstinence figures. Of those subjects who reported abstinence at one year, 41 percent had relapsed during the follow-up period. Schachter[13] has noted that cumulative success rates in self-initiated quitting substantially exceed commonly reported clinical findings for single quit attempts.

Perhaps the spontaneous initiation of renewed quit attempts observed in the present study can be facilitated through systematic intervention. Little research has been devoted to recycling individuals who fail to sustain abstinence. Too much emphasis may have been devoted to one-shot (albeit relatively intensive) treatment approaches. The need for recycling is underscored by the fact that outcomes in our more recent studies have been generally less successful than was true for studies that we undertook five to 10 years ago. This is true despite the fact that smoking history and demographic characteristics of our clinic enrollees have remained remarkably similar over time. More emphasis should be placed on viewing cessation as a process rather than as a discrete event.[14] Given that most quit attempts end in failure, the importance of eliciting multiple attempts over extended periods of time becomes obvious.

APPENDIX

American Lung Association Clinics

The basic American Lung Association clinic format consists of an orientation session and seven additional 90 minute to two-hour sessions over a seven-week period. The first session provided an in-depth discussion of the general health effects of smoking. Emphasis was placed upon the fact that subjects can quit and that help and encouragement to do so would be available in the program. The second session introduced coping strategies for confronting urges to smoke. By the third session, subjects were expected to make a personal commitment to quitting and to state this commitment publicly at the session. Quit Day occurred at this third session. At a fourth session held two days later benefits of quitting were reiterated and possible withdrawal symptoms were discussed.

The remaining sessions were focused upon maintenance and included a consideration of healthier, more enjoyable nonsmoking life-styles. Discussion included relaxation techniques, exercise or physical fitness programs, avoiding weight gain, and coping with tension. Session 7 was a celebration in which subjects elected options such as a wine and cheese party or dinner at a restaurant. This event was intended to emphasize enjoyment of their new life-style (and also to provide practice in confronting what for many might be a difficult situation). At this time awards were given to everyone who completed the program. Although the American Lung Association clinic has approximately one-half the sessions of the Lando clinic, total clinic contact hours for the two methods are approximately equal. Twelve of 31 American Lung Association clinics followed a modified format adopted by the American Lung Association as an alternative for worksite settings. This modified format entailed 10 one-hour sessions. No significant differences in outcome were found between the traditional and modified formats.

American Cancer Society Clinics

Treatment consisted of an orientation session plus four, one-hour group sessions over a two-week period. Emphasis was placed upon individualization of program content by the facilitator and eliciting the active involvement of group members. The focus of treatment was again upon coping strategies and a positive smoke-free life-style. However, instructions to clinic leaders placed relatively more weight upon individual situations as opposed to group process. Subjects were informed that smokers who take two weeks to quit smoking are as likely to be successful as are smokers who take two months to quit. Quitting is again described as a two-part process: 1) stopping, and 2) staying stopped. Unlike the other clinic methods there is no set target date for abstinence, although participants are expected to quit during the latter half of the program.

Laboratory Clinics

Treatment consisted of 16 sessions (approximately 45 minutes to one hour in length) over a nine-week period. The first three weeks were devoted to preparation for quitting and the final six weeks to maintenance. A specific preparation technique involved a nicotine fading procedure adapted from that originally described by Foxx and Brown.[15] Subjects switched brands on a 30–60–90 percent weekly reduction schedule. They were free to choose any cigarette that provided the required nicotine dosage. Subjects also underwent a smokeholding technique similar to that described in 1979 by Kopel, Suckerman, and Baksht[*] during two clinic sessions held the week prior to the quit date. Smokeholding consisted of two sets of 10 trials separated by a five-minute break.

Subjects attended eight maintenance sessions over a six-week period following the quit date. Considerable emphasis was placed upon relatively unstructured group discussion. In addition, subjects signed contracts calling for specific rewards for abstinence. Group discussion tended to emphasize problem-solving. Subjects were encouraged both to discuss their own problems and to suggest possible strategies for other group members. The laboratory clinics are described in greater detail elsewhere.[6,16]

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[*] Kopel S, Suckerman K, Baksht A: Smoke holding: an evaluation of physiological effects and treatment efficacy of a new nonhazardous aversive smoking procedure. Paper presented at a meeting of the Association for Advancement of Behavior Therapy, November 1979

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National Nursing Shortage Commission Announced

A 15-member Commission on the National Nursing Shortage has been established, as announced in the February 23 *Federal Register*. The Commission will advise federal officials—including the DHHS Secretary, the Assistant Secretary for Health, and the HRSA Administrator—on projects implementing the recommendations of the Secretary's Commission on Nursing. The projects should attempt to make optimal use of available resources from federal, state, and local government and private organizations. The recommended projects will focus on five areas:

- Recruitment and the educational pathway;
- Retention and career development;
- Restructuring nursing services and effective utilization of nursing personnel;
- Data collection and analysis requirements; and
- Information systems and related technology in nursing.

The Commission will consist of four ex-officio members and 11 members appointed by the Secretary. The ex-officio members will be:

- A Health Care Financing Administration representative;
- The Chief Nurse Officer of the Public Health Service;
- The Director of the Division of Nursing, HRSA; and
- The Director of the National Center for Nursing Research, National Institutes of Health.

The appointed members will be selected as follows:

- Four from the nursing community;
- Four from health care providers and other nurse employers;
- One from third party payers;
- One representing economics and data policy fields; and
- One from the general public.

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Benzene, Benzo(a)Pyrene, and Lead in Smoke from Tobacco Products other than Cigarettes

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Abstract: Benzene, benzo(a)pyrene (BaP), and lead in mainstream smoke from cigars, roll-your-own (RYO) cigarette and pipe tobaccos were sampled to evaluate their potential health significance. Results with reference cigarettes were consistent with published values, providing support for the methodology employed. The emissions of benzene and BaP, expressed as mass emitted per gram of tobacco consumed, were similar for all products evaluated; for benzene, the mean values for cigars, RYO cigarette and pipe tobaccos were 156 ± 52, 68 ± 11, and 242 ± 126 µg/g, respectively. Mean values for BaP were 42 ± 7 and 48 ± 4 ng/g for cigars and RYO cigarette tobacco, respectively. Lead values were below the limit of reliable quantitation in all cases. The mean benzene concentrations in a puff ranged from 1 to 2×10^5 µg/m^3 for cigars, RYO cigarette and pipe tobaccos. For BaP, the puff concentration averaged about 60 µg/m^3 for cigars and RYO cigarette tobacco. The results suggest that smoking cigars, pipes or RYO cigarettes leads to potential exposures which exceed the No Significant Risk levels of benzene and BaP set pursuant to California's Proposition 65. These tobacco products are now required to bear a health hazard warning when sold in California. We recommend that this be adopted as national policy. (*Am J Public Health* 1990; 80:560–564.)

Introduction

Cigarettes sold in the US are required by federal law to bear warnings of their health hazard. Other tobacco products, viz., cigars, pipe, and roll-your-own (RYO) cigarette tobaccos are not covered by federal law. A previous publication discussed California's Safe Drinking Water and Toxic Enforcement Act of 1986 (Proposition 65), including the requirement for warnings on commercial products containing or forming chemicals "known to the State" to cause cancer or other reproductive toxic effects.[1] Accordingly, the levels of three tobacco smoke constituents subject to Proposition 65 regulation, benzo(a)pyrene, (BaP), benzene, and lead, were determined. For benzo(a)pyrene, there is limited evidence of carcinogenicity in humans, but sufficient evidence of carcinogenicity in experimental animals.[2] Benzene has been shown to increase the incidence of leukemia in exposed workers, and to produce cancer in animals at various sites by several routes of exposure.[3] The reproductive toxicity of lead has been well documented in both human and animal investigations.[4]

Tobacco smoke includes both particle-phase and gaseous combustion products.[5,6] The particle-phase material represent about 8 percent (by weight) of the total. The gas-phase constituents include air (58 percent by weight), excess N_2 (15 percent by weight) and vapor constituents (19 percent by weight), the latter being hydrocarbons, oxygenated organic compounds and other gases (e.g., carbon monoxide). Polyaromatic hydrocarbons (PAHs) are found in both phases, the low molecular weight materials (e.g., naphthalene) being in gas phase, together with single ring aromatic hydrocarbons (e.g., benzene), and the five-ring PAH, benzo(a)pyrene (BaP) being almost exclusively in particle phase under usual conditions. Lead in cigarette tobacco smoke exists primarily in the particle phase. Commercial cigarettes from five countries showed a range for lead in mainstream tobacco smoke of 0.4 to 1.0 µg/cigarette in the particle phase, compared with 0.1 to 0.2 µg/cigarette in the gas phase.[7]

Mainstream cigarette tobacco smoke samples are routinely generated with laboratory "smoking machines" using Federal Trade Commission (FTC) standard smoking conditions: a two-second puff of 35 ± 2 ml volume, once/min, and a butt length of 23 mm.[8] FTC guidelines are lacking for RYO cigarette and pipe tobaccos, as well as for cigars. The Tobacco Research Council of London (TRC) specified for cigars the same parameters as used with cigarettes, with a butt length of 25 mm.[9] A cigar smoking protocol using a 1.5 sec puff, 20 ml puff volume, 40 sec puff interval and 33 mm butt length has also been recommended.[10] No standard protocol is available for testing pipe tobacco or RYO cigarettes, to our knowledge.

The present study measured and compared cigars, pipe tobacco and RYO cigarettes, for the levels of benzene, BaP, and lead in their mainstream smoke; and compared the mainstream smoke results to the daily exposure levels of benzene, BaP and lead judged in California Proposition 65 to pose no significant risk. Because of time limitations, pipe tobacco smoke was assessed for benzene levels only.

Methods

Tobacco Products

The tobacco products evaluated are described in the Appendix. One cigar was smoked for each lead or BaP plus benzene sample. The RYO cigarettes were prepared with a rolling machine intended for private use, with a rice paper wrapper ("Abadie" brand), length 70 mm, to which was added a pre-weighed amount of tobacco. Three such cigarettes were smoked per sample. Pipe tobacco was evaluated when smoked in a pipe, as well as in RYO cigarettes, prepared as described above, weighing 1.50 ± 0.02 g, smoking one cigarette per sample.

For pipe smoking, 1 to 2 g of tobacco were smoked in a pre-conditioned wooden pipe (bowl dimensions 20 mm I.D., depth at center 39 mm, stem length 60 mm from center of bowl) requiring 3 to 4 g of tobacco for filling as follows: the pipe was filled loosely with tobacco to the top of the bowl, the tobacco compressed in half, the bowl refilled and the tobacco

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compressed to about three-fourths of a full bowl. The tobacco consumed was approximated from the weight difference for the pipe plus tobacco before and after smoking.

Sample Collection and Smoking Parameters

The smoking parameters adopted here for cigars, pipe and RYO cigarette tobaccos retained the puff duration and volume of the FTC guidelines and varied the puff frequency as necessary to keep the tobacco burning and (for cigars) to provide a smoking time <1 hour. Pipe tobaccos were evaluated as RYO cigarettes as well as when smoked in a pipe, to assess whether smoking conditions played a major role in determining emissions of the pollutants measured.

A smoking system (Figure 1) drew air through the cigar, cigarette or pipe for brief periods to simulate a smoker's puff. The system was adjusted to a 35 ± 2 ml puff volume per 2.00 ± 0.05 sec puff using a bubble flowmeter attached upstream of the system. The RYO cigarettes prepared from cigarette tobacco followed FTC smoking parameters (i.e., two sec puff, 35 ml puff volume, once/min, butt length 23 mm). The RYO cigarettes prepared from pipe tobacco employed the same parameters except for puff frequency, which was increased to two/min. Cigar testing followed TRC-recommended parameters (two sec duration, puff volume 35 ml, butt length 25 mm),[9] with puff frequency altered to two/min. With pipe tobacco smoked in a pipe, a puff frequency of 12/min was used. Before smoking, tobacco products were equilibrated at room temperature for 24 to 48 hours in a chamber at about 60 percent relative humidity.

The sampling scheme employed a Gelman A/E glass fiber filter to collect BaP, or an 0.8 μm pore size (Millipore AA) cellulose ester membrane filter to collect particle-phase lead. An 8 × 110 mm glass tube filled with granular, activated charcoal (400 and 200 mg charcoal in the front and rear sections, respectively), downstream of the glass fiber filter, collected benzene with a measured front section collection efficiency >99.99%. The 47-mm diameter filters were contained in a Teflon filter holder. To minimize filter plugging with smoke constituents from cigars, filters were replaced when half the cigar had been consumed. By this strategy, the decrease in flow rate during filter loading was ≤20 percent. The two filters were extracted together to recover the BaP or lead. Single filter collection with ≤20 percent decrease in flow rate was achieved with the other tobacco products. The decrease in flow rate caused an estimated ≤10 percent positive error in calculating the total air volume sampled, and a corresponding negative error in calculated concentrations of benzene and BaP in a puff.

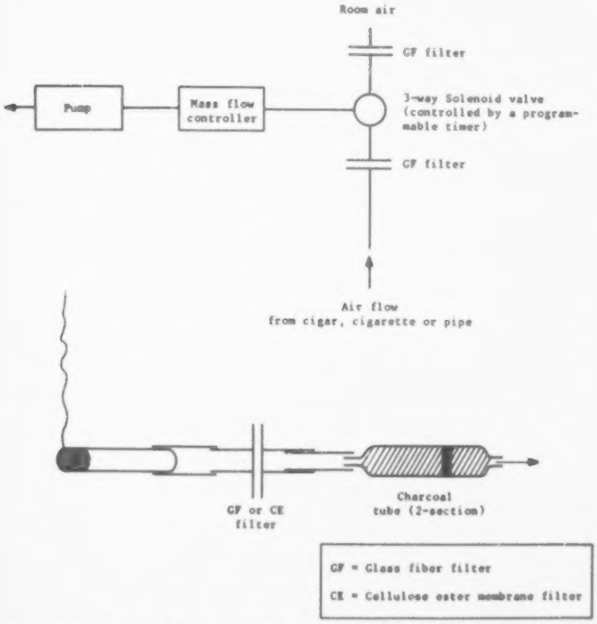
Room air
GF filter
Pump
Mass flow controller
3-way Solenoid valve (controlled by a programmable timer)
GF filter
Air flow from cigar, cigarette or pipe
GF or CE filter
Charcoal tube (2-section)
GF = Glass fiber filter
CE = Cellulose ester membrane filter

FIGURE 1—Schematic Drawing of the Collection of Tobacco Smoke Particulate and Vapor Phase Constituents

Chemical Analysis

BaP was extracted by one-hour ultrasonic agitation of the filter(s) in 70:20:10 v/v/v dichloromethane-methanol-benzene. Following removal of acidic and alkaline constituents by liquid-liquid extraction, and treatment with silica gel, samples were analyzed for BaP by reverse phase, high performance liquid chromatography, with fluorescence detection. Using NBS Standard Reference Material 1649 (urban dust), the recovery of BaP from the sample cleanup and analysis averaged 77 ± 4 percent. Benzene was extracted from the charcoal-filled collection tubes with CS_2 and directly analyzed by gas chromatography with photoionization detection.[11] Independent confirmation of the identity of the GC peak as benzene, as well as an interlaboratory comparison of benzene quantitation, was provided by gas chromatography-mass spectrometry (GC-MS) analysis on three CS_2 sample extracts. Benzene results from the two laboratories for the three extracts agreed, on average, within 10 percent. Lead was extracted from filter samples in concentrated HNO_3, and analyzed by flameless atomic absorption spectroscopy, using a heated graphite analyzer.[12] Blank trials with unlit tobacco showed no significant levels of BaP, benzene or lead.

Results are reported per gram of tobacco consumed (i.e., the total weight of tobacco consumed during as well as between puffs) as well as per cigar, or cigarette. Tobacco products were also compared based on average concentrations in the air stream sampled (i.e., puff concentrations). Comparison of results among the tobacco products is complicated by differences in puff frequencies. For example, at two puffs/min, a greater proportion of the tobacco will be consumed in generating mainstream smoke, compared to one puff/min. It is inappropriate to normalize results to a common puff frequency (e.g., by dividing results obtained at two puffs/min by two), since increased puff frequencies cause higher burning temperatures, enhance pyrolysis and decrease the amount of combustion or pyrolysis products retained on unburned tobacco in the butt.

Results

Quality Assurance

A Kentucky reference cigarette, type 2R1 (University of Kentucky), was evaluated for benzene and BaP production in the mainstream smoke. The procedures followed were the same as those for RYO cigarettes with cigarette tobacco; three cigarettes (average cigarette weight 1.18 g) were used to load a single charcoal tube. The benzene results for two such charcoal tube samples were 56 ± 2 and 67 ± 6 μg/cigarette (67 ± 2 and 81 ± 6 μg/g of tobacco burned), respectively. These results compare to 75 μg/cigarette reported for a similar Kentucky reference cigarette 1R1,[13] using a commercial smoking machine with the same smoking parameters. Smoke constituent deliveries for the two reference cigarettes are expected to be equal within about 30 percent.* Our benzene results for reference cigarette 2R1 averaged 18 percent lower than the published results for reference cigarette 1R1. Mainstream smoke BaP results for two trials, corrected for 77 percent recovery, were both 32 ng/cigarette (39 ng/g of tobacco burned). This compares to the 28.4 ± 3.4 ng/cigarette previously reported for this cigarette, using the same (FTC) smoking parameters.[14]

To assess the accuracy and precision of the analysis for lead at the relatively low levels expected, four blank cellulose ester membrane filters were each spiked with 0.1 µg lead from an aqueous solution of lead nitrate, and then subjected to extraction and analysis. The recovery averaged 83 ± 19 percent (n = 4).

Benzene

Mean tobacco smoke results, by brand, are shown for cigars and RYO cigarette tobacco in Table 1. Expressed by amount per cigar or cigarette, the benzene values for cigars and cigarettes showed a wide range of values. Brand F cigars were the only filter-tipped tobacco product evaluated in the present study. Benzene results for these filter-tipped cigars were similar to those for the others. The range in benzene results is reduced when expressed per gram of tobacco burned.

Benzene levels from RYO cigarettes were similar to those for the Kentucky reference cigarettes, smoked following the same parameters, and, expressed per gram of tobacco burned, averaged about a factor of two below those for cigars. This difference reflects, at least in part, the factor of two difference in puff frequency. Table 1 also shows average concentrations calculated for benzene in a puff, as well as the average number of puffs required with each product. The concentrations ranged from 0.7×10^5 to 3.3×10^5 µg/m^3, about 10^4 times those in urban ambient air.[15]

Table 2 summarizes pipe tobacco smoke benzene results for single cigarettes rolled from pipe tobacco, and for the same tobacco smoked in a pipe, expressed per gram of tobacco burned as well as by average puff concentrations. Again, ranges are wide. On average, results for pipe smoking were 2.5 times those in the RYO cigarette configuration. The difference probably reflects the much higher puff frequency used in pipe smoking, 12 vs two puffs/min, and the correspondingly greater proportion of the tobacco consumed in generating mainstream smoke. The mean result in the RYO cigarette configuration, 139 ± 45 µg/g, compares to 156 ± 52 µg/g for cigars, also smoked with two puffs/min. Thus, when tested using the same smoking parameters, pipe tobacco and cigars were not significantly different regarding mainstream smoke benzene levels. The puff concentrations from pipe tobacco smoked in both configurations are similar to those found for cigars and RYO cigarettes from cigarette tobacco.

BaP

The coefficient of variation (C.V.) for BaP measurement, with each brand of cigar or RYO cigarette tobacco evaluated in two or more trials, varied from 3 to 24 percent (Table 1). The BaP levels ranged from 96 to 292 ng per cigar, and were about 40 ng/cigarette for both RYO cigarette tobaccos tested. BaP data for Brand F filter-tipped cigars were not obtained. Expressed per g of tobacco burned, the range in BaP results is sharply compressed, and the results are similar for cigars and cigarettes. Both cigar and RYO cigarette BaP results are also similar to those obtained with the Kentucky reference cigarettes, 39 ng/g. Calculated puff concentrations of BaP showed a correspondingly small range for all tobacco products evaluated, and no difference between cigars and RYO cigarette tobacco, with an overall mean of 61 ± 14 µg/m^3 (C.V. 23 percent). It is unclear why differences in puff frequency between cigars and RYO cigarettes appear to influence benzene levels and not those for BaP. Benzo(a)pyrene concentrations are about 10^5 higher than those usually found in urban ambient air.[16]

Lead

Mainstream tobacco smoke particulate lead results for all cigars and RYO cigarette tobaccos tested were below the limit of reliable quantitation, 0.2 µg/cigar or cigarette.

Discussion

The yields of benzene and BaP, per gram of tobacco burned, in the non-cigarette tobacco products tested here are similar to those previously reported for a reference cigarette. However, the lead values obtained are lower than those in previous work (e.g., 0.4 to 1.0 µg Pb/cigarette cited in reference 7). This is consistent with a decrease in atmospheric lead observed in recent years, concomitant with the

TABLE 1—Benzene and BaP in Mainstream Tobacco Smoke from Cigars and Roll-Your-Own Cigarettes[a]

			Benzene		BaP		Puff Conc. (µg/m^3)		
Sample Type	Brand	N[b]	µg/cig[c]	µg/g	ng/cig[c]	ng/g	Benzene ($\times 10^{-5}$)	BaP	Mean No. Puffs
Cigar	A	7	1025 ± 57	151 ± 9	248 ± 44	36 ± 5	2.9 ± 0.3	63 ± 12	107
	B	1	977	165	206	35	1.9	41	137
	C	2	316 ± 17	92 ± 4	135 ± 1	39 ± 1	1.7	71	53
	D	2	482 ± 65	246 ± 31	96 ± 23	49 ± 12	3.3	84	43
	E	1	732	122	292	49	1.4	54	157
	F	2	228 ± 68	157 ± 59	ND[d]	ND[d]	2.3 ± 0.7	N.D.	26
RYO Cigarette[e]	A	2	61 ± 13	75 ± 16	37 ± 3	45 ± 4	0.9 ± 0.1	53 ± 1	65
	B	2	50 ± 10	60 ± 13	43 ± 4	51 ± 3	0.7 ± 0.1	63 ± 4	58
Mean for Cigars				156 ± 52		42 ± 7	2.3 ± 0.7	63 ± 16	
Mean for Cigarettes				68 ± 11		48 ± 4	0.8 ± 0.1	58 ± 7	

a) Results are means and standard deviations.
b) N = number of trials.
c) "Cig" refers to both cigarettes and cigars.
d) ND = not determined.
e) Roll-your-own cigarettes prepared from cigarette tobacco.

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*Jenkins RA: Oak Ridge National Laboratory, personal communication, Sept. 5, 1988.

TABLE 2—Benzene in Mainstream Smoke from Pipe Tobacco[a]

Brand	Benzene (μg/g)		Benzene Puff Conc. (μg/m³) × 10⁻⁵		Mean No. Puffs[c]
	Cig. Config.[b]	Pipe Config.	Cig. Config.[b]	Pipe Config.	
A	85 ± 5	253 ± 25	0.54 ± 0.01	1.3 ± 0.2	45, 72
B	114 ± 23	277 ± 20	0.9 ± 0.3	1.50 ± 0.04	37, 84
C	144 ± 48	354 ± 34	1.1 ± 0.5	1.9 ± 0.1	38, 90
D	145 ± 11	363 ± 2	1.1 ± 0.1	1.51 ± 0.01	40, 72
E	207 ± 12	473 ± 22	1.5 ± 0.1	2.6 ± 0.6	40, 78
Mean value	139 ± 45	344 ± 86	1.0 ± 0.4	1.8 ± 0.5	
Overall mean	242 ± 126		1.4 ± 0.6		

a) Results for each brand are the mean and standard deviation for two trials.
b) "Cig." indicates "RYO cigarette."
c) The two values refer to RYO cigarette and pipe configurations, respectively.

decreased use of leaded gasoline. For example,[17] over the period 1975 to 1984, the highest calendar quarter mean lead concentration decreased at a downtown Los Angeles site from 4.9 to 0.7 μg/m³. Similar decreases have been observed in North Carolina, an important tobacco-growing State.**

Under regulations implementing California's Proposition 65,[18] carcinogens are deemed to pose no significant risk (NSR) if daily exposures do not exceed those expected to produce a lifetime excess cancer risk of 1 in 100,000. The daily NSR level set for benzene is 20 μg, and for BaP, it is 0.06 μg.

The significance of the levels of these chemicals measured in mainstream tobacco smoke can be judged by comparison to the NSR levels. For example, the amounts of benzene in mainstream smoke from burning one 1.2-g RYO cigarette, a 7-g, 16-cm ("Churchill size") cigar (e.g., cigar brand E), or a bowlful (e.g., 3.5 g) of pipe tobacco were approximately 3, 36 and 60 times the NSR level, respectively. The corresponding amounts of BaP from a RYO cigarette and the Churchill-size cigar were approximately 0.9 and five times the NSR level, respectively. Thus, benzene and BaP in mainstream tobacco smoke from less than one Churchill-size cigar, two RYO cigarettes or a bowl of pipe tobacco exceeded daily NSR levels under California's Proposition 65, using simulated smoking protocols. With smokers, actual exposures to these chemicals will vary by the degree to which the smoke is inhaled. Similarly, the puff concentrations reported are not intended to reflect concentrations potentially inhaled by smokers, since the puff is subject to substantial dilution by air not drawn in through the cigarette.

California's interest in limiting toxic exposures, as expressed through Proposition 65, may portend an increased national focus on the adverse health consequences of smoking. Recently, manufacturers of cigars, RYO cigarette and pipe tobaccos have added warning labels for these products sold in California. We recommend that warning labels be made a standard requirement for all such products sold in the US.

ACKNOWLEDGMENTS

Tobacco samples were obtained from retail stores by the Food and Drug Branch, California Department of Health Services. We wish to acknowledge the helpful comments of Dr. Roger Jenkins, Oak Ridge National Laboratory, regarding methodology for generating and sampling tobacco smoke. Dr. Jenkins also supplied the Kentucky reference cigarettes. We also acknowledge the helpful comments of Dr. H.C. Pillsbury, Federal Trade Commission. GC-MS determinations of benzene were provided by Dr. J. Hsu.

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APPENDIX
Tobacco Products Evaluated and Smoking Parameters Used[a]

Product Type	Brand	Weight Range, g	Length, cm	Puff/min
Cigar	A	7.1–8.6	13.8	2
Cigar	B	7.1–7.3	16	2
Cigar	C	4.4–4.8	11.5	2
Cigar	D	2.5–2.8	10.4	2
Cigar	E	6.4–7.1	15.9	2
Filter-Tipped Cigar	F	1.7–2.0	10.0	2
RYO[b] cigarette tobacco	A	1.2 ± 0.05	7.0	1
RYO cigarette tobacco	B	1.2 ± 0.05	7.0	1
Pipe tobacco	A	—[c]	—[c]	2,12[d]
Pipe tobacco	B	—	—	2,12
Pipe tobacco	C	—	—	2,12
Pipe tobacco	D	—	—	2,12
Pipe tobacco	E	—	—	2,12

a) Puff volumes were 35 ± 2 ml, and puff duration were 2.05 ± 0.05 sec in all cases. Butt lengths for cigars and RYO cigarettes were 25 and 23 mm, respectively. All tobacco products were preconditioned at ca. 60% R.H. for 24 to 48 hr before smoking. All uncertainty values in this table reflect ranges, not standard deviations.

b) Roll-your-own.

c) Evaluated as RYO cigarettes, 1.50 ± 0.05 g and 7.0 cm length each, and when smoked a pipe, consuming 1 to 2 g of the 3 to 4 g of tobacco used in filling the bowl.

d) Two/min in the RYO cigarette configuration, 12/min in the pipe.

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Rockefeller Foundation Grants: Global Environmental Program Announced

The Rockefeller Foundation has announced its new Global Environmental Program, a grant-making initiative to "promote the skills, attitudes, relationships and institutions necessary for environmentally sound international development." Whether humans can live in balance with the natural environment and whether all the countries and cultures of the world can achieve a reasonable and just standard of living are two emerging aspects of the same issue which will be addressed by the new program.

The new Global Environmental Program is part of the Foundation's International Program to Support Science-Based Development built on the premise that the well-being of people in developing nations can be advanced in important ways through the application of science and technology. The environmental initiative takes into account, however, that the application of science and technology to development in the industrialized nations has not always been kind to the environment. The challenge before the world community is to promote a more equitable sharing of the world's resources and yet avoid repetition of environmentally unsound development practices.

The Global Environmental Program will begin with four broad areas of interest for grant and fellowship activity:

- supporting young environmental scientists, analysts, policymakers, and community leaders who can be architects of future policy in development and the environment;
- advancing the development of economic analysis which can guide and illuminate the process of environmentally sound development;
- laying the groundwork for multilateral bargains and accords that will be necessary to address selected environmental issues; and
- Seeking innovative ways to help the United States public prepare for new environmental realities.

No special form is required for grant applications. However, they should include:

- a description of the project, with clearly stated plans and objectives;
- a comprehensive plan for the total funding of the project during and, where applicable, after the proposed grant period; and
- a listing of the applicant's qualifications and accomplishments and, if applicable, a description of the institutional setting.

Address grant applications and requests for further information to:

Global Environmental Program
The Rockefeller Foundation
1133 Avenue of the Americas
New York, NY 10036, USA

Long-Term Effectiveness of Mass Media Led Antismoking Campaigns in Australia

JOHN P. PIERCE, PHD, PETRA MACASKILL, MAPPSTAT, AND DAVID HILL, PHD

Abstract: A community antismoking campaign began in Sydney, Australia in 1983, and in Melbourne in 1984. These campaigns purchased prime-time television advertising spots to set the community agenda. An intense effort was made to ensure that antismoking activities were maximized at the school, organizational, and community level. Smoking prevalences in both cities from 1981 were fitted with a statistical model to identify any underlying trend, to assess any immediate impact, and to assess the longer term effect of continuing to conduct such campaigns, i.e. to identify any change in the underlying trend. During the years before the antismoking campaigns, there was no observable trend in smoking prevalence in either city. At the beginning of the campaigns, there was an immediate drop of more than two percentage points in male and female smoking prevalence in both cities. Thereafter, a decline of about 1.5 percentage points per year was observed among males. No post campaign trend was observed in smoking prevalence for women in either city. These data support conducting coordinated community campaigns to reduce current smoking prevalence. (*Am J Public Health* 1990; 80:565–569.)

Introduction

Smoking prevalence in Australia did not decline much from 1974 to 1983 according to age standardized comparisons,[1,2] in contrast to the pattern seen in other Western developed countries.[3] For this reason, a number of individual states in Australia introduced a coordinated campaign approach to promoting change in health behavior, as pioneered in cardiovascular disease by the Stanford Three Community Project[4] and the North Karelia Project.[5]

In 1983 in Sydney and in 1984 in Melbourne, the State Health Departments established anti-smoking steering committees composed of personnel from the health departments, the major voluntary organizations, health professional societies, and university public health departments. Television commercials were developed to motivate smokers to quit smoking and to set the agenda for professionals whom the campaign team hoped to involve in promoting nonsmoking, given the evidence of the effectiveness of medical advice.[6-8] The effectiveness of school education programs[9] also made establishing such programs a high priority within the community. Although the campaign committees in both cities shared campaign materials and ideas, the actual activities undertaken within each city were dictated by the perceived local needs and opportunities.

A detailed evaluation was undertaken of the first year of the campaign in Sydney using Melbourne, which did not have a campaign that year, as a reference community.[10,11] More than 80 percent of the population of Sydney remembered the television commercials. After these commercials were aired, use of services such as a telephone quit line and enrollment in cessation classes increased. The campaign was also associated with higher levels of intention to quit in Sydney, a variable that had previously been shown to be associated with actual quitting.[12] Changes in smoking prevalence during the first campaign year were assessed by self-reported smoking behavior (after validation with salivary cotinine[13]) from both cross-sectional and longitudinal surveys. No decrease was seen in the reference city, Melbourne. The pooled estimate of the difference in smoking prevalence attributable to the campaign was 2.8 percent (95% confidence interval = 0.5, 5.1).[11]

Since 1984, the campaigns have continued in both Melbourne and Sydney, spearheaded each year by commercials shown on prime time television for six to eight weeks during the winter months. In this paper, we assess the long-term effect of this approach to reducing the prevalence of smoking. We used data from both cities to determine whether these key intermediate goals were achieved:

- High public awareness of the campaign during each campaign year;
- A strengthening of the stand of health professionals, particularly physicians, against smoking;
- An increased proportion of the smoking population who both believed that smoking was harming their own health and also were being influenced by their social network to quit smoking.

The effectiveness of the campaign was assessed by using self-reported smoking data from 1981 to 1987 for each city.

Methods

General Design

This project has a before and after design in each of two intervention cities with a temporal lag of 12 months between the onset of the intervention in the second city. Prevalence data were obtained for several years before the onset of the intervention, as well as for the duration of the campaign. In 1983 and 1986, more detailed questions about health beliefs, intentions, and social influences were asked in each city. Small random surveys were undertaken approximately two months after the completion of each component of the television campaign to assess recall of commercials.

Mass Media Used in Campaigns

Television—In both Sydney and Melbourne, television advertising made up about two-thirds of the media budget for the campaigns. Typically, commercial spots were purchased during prime or fringe time for approximately four weeks out of eight at the start of the campaign year. Approximately 40 spots per week were shown during a schedule designed to ensure at least 50 percent market penetration in the Sydney

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and Melbourne metropolitan areas during each year of the campaign.

Production of commercials followed research on messages that were likely to be effective among the target audience. These commercials were designed to use strong visual images of the health consequences of smoking. Further, they sought an emotive response among the audience to ensure attention and to maximize the likelihood that viewers would reassess the hazards of smoking and then possibly quit in the near future (for description, see Appendix).[12,15] All of the commercials ended with a "Quit line" telephone number that people could call 24 hours a day to hear a recorded informative and encouraging message about quitting smoking.

Other Mass Media—A concerted effort was made to coordinate all mass media advertising around a central theme that was dictated by the message in the major television commercial being run at the time. For billboard advertising, a simple message from the television commercial was displayed at numerous sites, especially those near mass transit locations. Advertising in newspapers included the normal large advertisement and, in Sydney, a section was purchased for journalistic coverage of smoking-related events and issues. Antismoking skits by major personalities complemented the normal quit smoking commercials on the radio. In addition, the campaign generated many news releases and events that resulted in substantial news coverage in all the mass media.

Physician's Offices

In both cities, there was active collaboration with the Australian Medical Association to develop and distribute printed resources that physicians could use to encourage and help their patients quit smoking. In Melbourne, a special training videotape about smoking was produced and shown widely to family physicians. In Sydney, a special training program was available to family physicians to enable them to more effectively counsel their patients to stop smoking.[16]

Schools

In Sydney, two school programs were developed: a comprehensive kindergarten through 12th grade curriculum and a grade-specific peer or teacher led curriculum. By 1986, 80 percent of all schools in Sydney had at least one person who was trained by the campaign team in conducting the program of choice for that school. Sixty-four percent (88 percent of the 73 percent who responded) of these schools reported using these materials during the 1986 academic year. Further, the campaign team sponsored theatre performances, rock concerts, and sports activities with antismoking themes.

In Melbourne, kindergarten through 12th grade curriculum resources on smoking were made available to all schools, and project workers encouraged and helped teachers to use these resources. By 1986, in sample surveys of school children in both Sydney and Melbourne, more than 70 percent of 12–15-year-olds reported receiving at least minimal health education about smoking during the year.[17]

Other Community Activities

Each year, the campaigns either organized smoking cessation clinics or informed the population in each city about the availability of clinics. Training programs were available for health educators who might be appointed to local government or other affiliated positions. In addition, quit smoking display stands were regularly erected and staffed in shopping malls and at public fairs.

Assessing Smoking Prevalence

Between 1981 and 1987, 68,136 males and 70,634 females aged 16 years and over from the suburbs of metropolitan Sydney and Melbourne were interviewed in their homes as part of an ongoing weekly survey. Sampling procedures, reported in detail elsewhere,[11] involved selecting a household starting point at random from the electoral roll (in Australia, it is compulsory for citizens over 17 years of age to be registered on the electoral roll). Field interviewers proceeded to move from house to house in a clockwise direction until they obtained a cluster of 10 interviews from separate households. Within the household, the youngest person aged 16 years and over of a selected sex was interviewed. The resulting sample has been shown to be representative of the population on major sociodemographic variables.[11]

A smoker was defined as anyone who responded positively to questions on whether he or she smoked factory-made cigarettes, roll-your-own cigarettes, cigars, or a pipe. In two separate subsamples, self-reported smoking was validated with biochemical analysis of saliva for the presence of cotinine.[13]

Measuring Campaign Exposure

For each year of the campaign, recall of the television commercials was assessed on separate random samples of the population of at least 1,000 persons in the three months following the completion of the television phase of the campaign. Shortly after the completion of each annual television campaign, a subsample of persons in Sydney were shown a set of six still photographs taken from the commercials. They were first asked whether they had seen the commercial and, if they responded yes, what the message was (care was taken that no "giveaway" words or cigarettes appeared in the photographs). In Melbourne, the respondents were asked whether they had seen any television commercials against smoking, recently. The interviewer then probed all positive responses and recorded details of the message recalled.

Assessing Predictors of Quitting

In 1983 and 1986, additional random samples of the population of Sydney and Melbourne were surveyed. The Sydney survey included 271 smokers in 1983 and 557 smokers in 1986. The Melbourne sample included 217 smokers in 1983 and 550 smokers in 1986. Smoking respondents were asked to agree or disagree with a set of statements previously shown to be scales representing health beliefs and social influences.[12] The items in the health belief scale were "Smoking cigarettes doesn't greatly affect my health," "The problems with smoking only affect heavy smokers," and "I'd rather enjoy life as a smoker than live a little longer without cigarettes."

The items in the social influence scale were "I find that my smoking annoys people around me who don't smoke," "What I dislike about smoking is the smell that it leaves on clothes", "All the physicians that I know are strongly against smoking," "My close friends would prefer that I didn't smoke," and "I find that people are talking a lot more about smoking and the problems of giving up." Internal consistency reliability (Cronbach's alpha) on these scales have been previously reported at 0.57 on the health benefit scale and 0.56 on the social influence scale.[12]

Statistical Analysis of Prevalence

For the purpose of analysis, data were grouped into intervals of six months for each city and were adjusted for age and sex to the estimated population of Sydney in 1986 using direct standardization. Salonen, *et al*,[18] have suggested using multiple linear regression models to evaluate the effectiveness of community-based intervention programs by testing for interaction effects between the type of community (intervention or reference) and the timing of the onset of the intervention effect. In this study, which lacked a nonintervention community, the following simplified form of the suggested model was fitted to each of the data sets.

$$P = \beta_0 + \beta_1 C + \beta_2 T + \beta_3 CT + \beta_4 S + \text{error}$$

where P represents the prevalence of smoking,

C is a binary indicator variable representing whether the data are from the precampaign (C = 0) or the postcampaign (C = 1) period,
T is the trend variable for time,
S is a binary indicator of season with a value of 1 for the first six months and −1 for the second six months of each year.

Using this model, β_0 is an estimate of the expected prevalence at the start of the campaign (T = 0) given the underlying trend in the precampaign period. β_1 is an estimate of the immediate effect of the intervention (as such an effect was demonstrated in a previous paper[15]). β_2 estimates the underlying trend in prevalence over the whole time period, and β_3 estimates the change in this underlying trend associated with the commencement of the campaigns. The coefficient β_4 provides an estimate of the effect of seasonal variation on the data. For each analysis, a model which included all the explanatory variables described above (full model) was fitted. A parsimonious model (adjusted for seasonal effect) was then developed by using backward elimination to remove variables, other than S, with the chosen level of significance set at 0.05. This reduced model identifies the pattern of change in prevalence in relation to the timing of the campaigns. A series of curvilinear (polynomial) models were also fitted to the data. Quadratic and cubic models were compared with the linear parsimonious models described above to assess whether the more complex models fit the data better.

Results

Population Knowledge of the Campaign

We assessed awareness about the antismoking campaign by asking people in each city what they recalled about the most public aspect of the campaign, the television commercials (Table 1). For all the major commercials in these campaigns, there was little difference between the recall of the commercial and identification of an appropriate antismoking message. In both cities, more than two-thirds of the community knew that an antismoking campaign existed in the years that such campaigns were conducted.

Action by Physicians

During the campaign, physicians' roles were assessed by asking respondents how strongly the physicians were against smoking; these data were then divided by age and sex for each city in 1983 and 1986 (Figure 1). Before the campaign, only 15 percent of males and females smokers over age 40 in Sydney thought their physicians were strongly against smoking. By 1986, this had increased dramatically to 47 percent for males and 50 percent for females. In those under the age of 40 in Sydney, this proportion increased from 36 percent to 48 percent in males and from 28 percent to 46 percent in females. In Melbourne males, this proportion increased from 35 percent to 41 percent in those under 40 years and from 28 percent to 35 percent in those over 40. In Melbourne females, this proportion increased from 27 percent to 47 percent in those under 40 years and from 33 percent to 37 percent in those over 40 years.

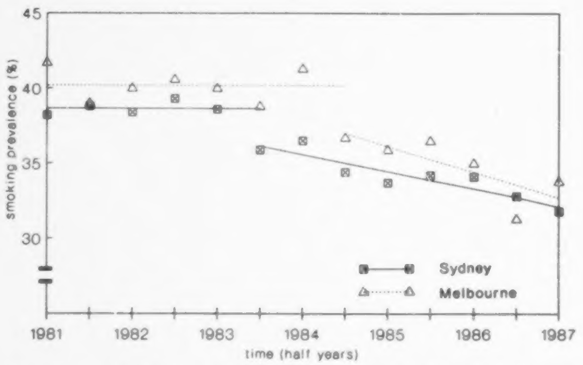

FIGURE 1—Age-Standardized Smoking Prevalence for Males in Sydney and Melbourne, 1981 to 1987

The Health Belief/Social Influence Variable

A previous study[12] had demonstrated that only a combination of personal beliefs about smoking and the environmental influences to quit was associated with an increased likelihood to quit. This proportion reporting both social influence and health beliefs was low in those over the age of 40 years in Sydney in 1983 (males, 10 percent; females, 12 percent). By 1986, this proportion had increased to 29 percent in males and 27 percent in females. In those under 40 in Sydney, males increased from 34 percent to 53 percent and females increased from 30 percent to 44 percent. In those over 40 years in Melbourne, males increased from 21 percent to 29 percent and females increased from 22 percent to 32 percent. In those under 40 years in Melbourne, males increased from 30 percent to 43 percent and females increased from 31 percent to 48 percent.

Smoking Prevalence

We considered both the full and parsimonious statistical models fitted to the age-standardized data for men and

TABLE 1—Percentage of Males and Females Who Recalled Television Advertising in Sydney and Melbourne

Year	% Males	% Females	N
Sydney (aided by picture prompt)			
1983	84	84	2,970
1984	72	80	1,046
1985*	79	76	996
1986	92	94	457
Melbourne (unaided)			
1984	70	67	886
1985	73	68	1,818
1986	68	67	1,830

*Only 14-to-19 year-olds.

women separately for each city. For men, these models fit the data very well, explaining 96 percent of the variation in Sydney and 91 percent in Melbourne. For females, the best models did not fit quite so well, explaining 72 percent of the variation in the Sydney sample and 79 percent in the Melbourne sample. However, the most parsimonious model for the Melbourne women sample was not consistent with the other models. A consistent model fitted almost as well and explained 68 percent of the variation.

Figures 1 and 2 are graphical representations of the trends identified in the parsimonious models in conjunction with the observed data. For this presentation, the consistent model was chosen for Melbourne women.

Thus, the best estimate of smoking prevalence among Sydney men (Table 2) indicated that prevalence did not change in the three years before the campaign (June 1983) and was 38.7 percent (β_0 in the parsimonious model). From 1981 to 1987, there was no significant underlying trend in smoking prevalence other than that associated with the campaign. The introduction of the campaign coincided with an estimated immediate drop in smoking prevalence of 2.5 percentage points in that six-month period. Continuation of the campaign was associated with a further drop of 1.12 percentage points per year ($2 \cdot \hat{\beta}_3$). These estimates are consistent with those provided by the full model.

For women, smoking prevalence had again been constant during the years before the campaign and was 31.6 percent in Sydney and 29.8 percent in Melbourne. In Sydney, there was an immediate drop of 2.6 percentage points associated with the introduction of the campaign but smoking prevalence did not significantly decline in the later years of the campaign. Again, this estimate is consistent with that from the saturated model.

Men in Melbourne responded similarly to those in Sydney. As Table 3 shows, the expected prevalence of smoking at the beginning of the campaign (June 1984) was 40.1 percent ($\hat{\beta}_0$). Again, there was no underlying trend. With the start of the campaign, there was an immediate drop of 2.9 percentage points. A further decline of 1.9 percentage points per year occurred in conjunction with the continued campaign. Once again these estimates are consistent with those from the full model.

In Melbourne, two models fitted the data and, as indicated earlier, model 2 is the preferred model. Smoking prevalence prior to the campaign was 30.9 percent. The immediate effect of the campaign was a 2.5 percentage point drop in smoking prevalence, again with no subsequent decline associated with the continued campaign. However, in this instance, the estimate from the full model is very different, highlighting the uncertainty that exists as to the true underlying model.

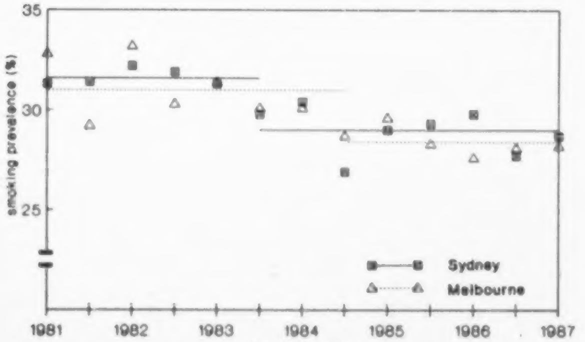

FIGURE 2—Age-Standardized Smoking Prevalence for Females in Sydney and Melbourne, 1981 to 1987

TABLE 2—Statistical Models for Smoking Among Males and Females in Sydney

		Full Model Estimate (SE)	Parsimonious Model Estimate (SE)
Sydney Males	$\hat{\beta}_0$	39.06 (0.67)	38.67 (0.28)
	β_1	−2.91 (0.79)	−2.52 (0.49)
	β_2	0.13 (0.20)	
	β_3	−0.69 (0.22)	−0.56 (0.10)
	β_4	−0.04 (0.18)	−0.04 (0.17)
		(R^2 = 0.96, 8df)	(R^2 = 0.96, 9df)
Females	$\hat{\beta}_0$	031.69 (1.02)	31.56 (0.42)
	β_1	−2.11 (1.20)	−2.61 (0.54)
	β_2	0.05 (0.31)	
	β_3	−0.23 (0.34)	
	β_4	0.38 (0.27)	0.32 (0.26)
		(R^2 = 0.77, 8df)	(R^2 = 0.72, 10df)

Comparison with other Models

The goodness of fit of these parsimonious linear models was compared with quadratic and cubic models. In each case, the best-fit linear model was a superior fit to either of the alternative models, except for Melbourne women, again emphasizing the problems in interpreting this set of data.

Discussion

This study suggests that smoking prevalence in the two major Australian capital cities did not decline prior to the campaign, as previously observed for the period 1974 to 1980 for Australia as a whole.[1-3] The mass media led antismoking campaigns coincided with a marked drop in smoking prevalence during the six-month period immediately following the start of the campaign for all persons in Sydney and for men in Melbourne (the effect on Melbourne women is less clear). The estimated size of this immediate effect was 2.6 percentage points in Sydney. A larger 2.9 percent effect of the antismoking campaign was seen among Melbourne men 12 months later. If we accept the model that suggests that women in Sydney and Melbourne reacted similarly to the antismoking campaign, the campaign immediately reduced smoking prevalence in Melbourne by 2.5 percent, an amount

TABLE 3—Statistical Models for Smoking Among Males and Females in Melbourne

		Full Model Estimate (SE)	Parsimonious Model Estimate (SE)	
Melbourne Males	$\hat{\beta}_0$	39.89 (1.00)		40.12 (0.42)
	$\hat{\beta}_1$	−2.64 (1.33)		−2.87 (0.93)
	$\hat{\beta}_2$	−0.06 (0.22)		
	$\hat{\beta}_3$	−0.89 (0.36)		−0.95 (0.27)
	β_4	0.58 (0.34)		0.57 (0.32)
		(R^2 = 0.91, 8df)		(R^2 = 0.91, 9df)
			Model 1	Model 2
Females	$\hat{\beta}_0$	29.77 (0.83)	29.40 (0.26)	30.93 (0.41)
	$\hat{\beta}_1$	−0.53 (1.10)		−2.52 (0.61)
	$\hat{\beta}_2$	−0.29 (0.19)	−0.37 (0.07)	
	$\hat{\beta}_3$	−0.04 (0.30)		
	β_4	0.64 (0.28)	0.65 (0.25)	0.56 (0.30)
		(R^2 = 0.80, 8df)	(R^2 = 0.79, 10df)	(R^2 = 0.68)

similar to the estimates of the effect obtained in both men and women for Sydney one year earlier. In the alternative model, women in Melbourne were the only group who recorded a decreasing prevalence of smoking during the precampaign period. Given that there was no downward trend in smoking prevalence throughout the 1970s,[1-3] this model must be considered less plausible.

For men in both cities, continued antismoking campaigns were associated with an average continued decrease in smoking prevalence of approximately 1.5 percent per year. The situation for women is less clear, but suggests that the continued campaigns in Sydney and Melbourne were not associated with any further decrease in smoking prevalence among women.

This analysis of changes in community smoking prevalence does not consider the effect that such campaigns may have on uptake of smoking. In any given year, the influence of new smokers between the ages of 16 and 20 years (the teen years included in this data set) on the overall prevalence will be small. Before we can conclude how effective these campaigns really are, changes in the uptake of smoking among teenagers need to be explored. However, even with the effect on uptake of smoking being unresolved, this study supports conducting coordinated mass media antismoking campaigns (using purchased television time) as an effective way to reduce smoking prevalence in the community. Whether conducting campaigns for up to four consecutive years is cost-effective will be the topic of a subsequent paper.

APPENDIX

TV Commercials by City and Year

	Year			
Commercial	1983	1984	1985	1986
"Sponge"[a]	S	S, M	M	
"I've had enough"[b]	S	S, M	M	
"Mrs. Holden"[c]		S		
"Yul Brynner"[d]				S, M
"Heartbeat"[e]			M	M
"Stairs"[f]				M
"Mirrors"[g]			S	
"Tough Guy"[h]			S	
"Mates"[i]			S	

S = Sydney, M = Melbourne.

[a] an analogy between a sponge and a lung soaking up tar.

[b] smoking histories of three people outlining reason for quitting included ashtray on a girl's face and modeling of calling the Quitline.

[c] a laryngectomy patient (with subtitles) outlining her smoking related problems with communications.

[d] a posthumous exhortation "Whatever you do, don't smoke."

[e] cardiograph used to outline smoking effects on circulation."

[f] mirror distortion to reduce fears of weight gain and stress with quitting.

[g] linked smoking to shortening a young girl's life.

[h] teens, visiting cancer patient, review smoking-related disease.

[i] outlined social problems teens might face if they smoke.

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A Twin Study of the Effects of the Vietnam Conflict on Alcohol Drinking Patterns

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Abstract: This study examines the association between military service in Southeast Asia and alcohol drinking patterns in 2,169 male-male monozygotic twin pairs who both served on active military duty during the Vietnam era (1965–75). Data on alcohol drinking were collected in 1987 by mail and telephone interview. The alcohol drinking measures include three indicators of abstention (lifetime abstainer, lifetime non-regular drinker, and current abstainer) and two indicators of consumption (average daily ethanol consumption and high consumption). In unadjusted and co-twin adjusted analyses, neither service in Southeast Asia nor combat exposure was significantly associated with any measure of abstention. In the co-twin adjusted analysis, there was no association of Southeast Asia service and combat exposure with average daily ethanol consumption. After adjustment for co-twin effects, 4.0 percent of non-Southeast Asia veterans were high consumers compared to 6.7 percent of Southeast Asia veterans who served in high combat. We conclude that prior military service in a war zone has a relatively modest long-term effect on the alcohol drinking patterns of male veterans. (*Am J Public Health* 1990; 80:570–574.)

Introduction

The relationship between service in the military and alcohol drinking patterns is a subject of considerable debate. Numerous clinical studies have concluded that veterans have high rates of alcoholism.[1–3] The evidence for increased alcohol consumption among veterans in non-clinical population samples is equivocal; some studies conclude that increased consumption among veterans is a function of demographic differences between veterans and nonveterans,[4,5] while other more recent research indicates that veterans consume more alcohol than nonveterans regardless of demographic differences.[6]

A number of investigations have examined the association between military service in Southeast Asia and alcohol consumption. Some studies have found that service in Southeast Asia and exposure to combat are associated with increased rates of alcoholism and problem drinking,[7,8] while other research has failed to demonstrate any association with alcohol consumption patterns.[9–12]

The potential for selection bias is a fundamental problem in all studies of alcohol drinking among veterans. Observed associations between military service experience and alcohol drinking patterns may simply reflect an underlying association of predispositional characteristics with both military service experience and alcohol drinking. The present study addresses this problem by examining alcohol drinking in monozygotic twin pairs contained in the Vietnam Era Twin Registry.

Methods

Data Sources and Study Eligibility

The Vietnam Era Twin Registry contains 7,375 male-male twins who both served in the military during the official Vietnam era (May 1965-August 1975). A complete description of the development of the registry can be found elsewhere,[13] and only a brief overview is presented here.

A survey of all twins on the registry was conducted in early 1987, with the goals of collecting information on a wide range of physical and psychological indicators of health. The overall case-wise response rate was 74 percent and the pair-wise response rate was 65 percent or 4,774 pairs. Twins were eligible for the present study if: both members of the pair responded to the survey, the pair was originally ascertained from the US Department of Defense computer files, and complete information was available for all alcohol drinking measures. After applying the exclusion criteria, the final number of eligible twin pairs was 4,048.

Zygosity was evaluated by using a series of questions on twin similarity and limited blood group typing obtained from the military records.[14] Of the 4,048 pairs meeting the eligibility requirements for the present study, 2,169 (53.6 percent) were monozygotic, 1,773 (43.8 percent) were dizygotic and 106 (2.6 percent) were of indeterminant zygosity. The present study focuses exclusively on the monozygotic twin pairs.

Service in Southeast Asia and Combat Exposure

Military service in Southeast Asia was determined by the response to the following question: "When you were in the military, were you stationed in Vietnam, Laos, or Cambodia; in the waters in or around these countries; or fly in missions over these areas?" Those men who responded in the affirmative were then asked a set of questions about 18 specific combat roles or experiences (see Appendix A). For each item the veterans indicated whether they had that role or experience while serving in Southeast Asia. An index of combat exposure was constructed by taking the summation of all positive responses for an individual. For analytical purposes the combat index was grouped into five levels: (non-Southeast Asia service; Southeast Asia service, no combat; low, combat index = 1–2; medium, combat index = 3–6; high, combat index > 6). Elsewhere,[15] we have demonstrated that in a subset of twins who completed the health survey twice more than a year apart, concordance was high (weighted kappa = 0.84) for the combat exposure index. Likewise, the combat index was found to be significantly correlated with objective indicators of threat to life and limb as measured by combat medals and decorations.

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Measures of Alcohol Drinking

A negative response to having had more than 20 alcoholic drinks in their entire life defines a lifetime abstainer. Veterans who were not lifetime abstainers, were asked at what age they started drinking alcoholic beverages regularly; as part of the response option to this question the veteran could indicate that he never drank regularly. This is referred to as lifetime non-regular drinker. The next question asked if the veteran drinks alcoholic beverages now. This provides a third measure of alcohol abstention which separates current drinkers from noncurrent drinkers. It should be noted that the lifetime abstainer measure is subsumed by the lifetime non-regular drinker measure, which is subsumed by the current abstainer measure.

For men who are current drinkers quantity was obtained as the average number of drinks consumed on days when the veteran had at least one drink; frequency was obtained in terms of the average number of days per week that the veteran had at least one drink. If the ethanol content of an unspecified drink is assumed to be 0.5 ounces, then average daily ethanol consumption can be calculated as: (frequency × quantity × 0.5)/7. Following the guidelines of the National Institute of Alcohol Abuse and Alcoholism,[16] average daily ethanol consumption can be categorized into four drinking levels: abstainers, average intake of 0.0 ounces ethanol daily; light drinkers, 0.01–0.21 ounces ethanol daily; moderate drinkers, 0.22–0.99 ounces ethanol daily; heavy drinkers, 1.00 or more ounces ethanol daily. A second measure of alcohol consumption is constructed based on the number of ounces of ethanol corresponding to the upper fifth percentile of the average daily ethanol consumption distribution. This measure, referred to as high consumption, conformed to a usual intake of five drinks per day (2.5 ounces of ethanol) in the present study.

Reliability of Alcohol Drinking Measures — In 1985 a pilot study was conducted of 150 veteran twin pairs born in the State of Connecticut. This pilot study was designed as a field test of the methodology to be used in the main study. The measures of alcohol drinking in the field test and the subsequent survey were identical. Because these 150 twin pairs were included in the main study using the Vietnam Era Twin Registry it is possible to evaluate the reliability of the five measures of alcohol drinking.

Complete alcohol drinking data were available from 191 individuals from both studies. Lifetime abstention had a kappa of 0.43, a Yule's Y[17] of 0.77, and a 97 percent overall agreement. Considerably less reliable was the lifetime non-regular drinker measure, which had a kappa of 0.10, Y of 0.37, and a 82 percent overall agreement. The current abstainer measure had a kappa of 0.44, a Y of 0.48, and an overall agreement of 78 percent.

Both of the alcohol consumption measures demonstrated similar levels of reliability. Average daily ethanol consumption had a weighted kappa of 0.55; overall agreement was 81 percent. The measure of high consumption had a kappa of 0.57, a Y of 0.72 and an overall agreement of 92 percent.

Statistical Analysis

The analysis begins by randomly selecting one member of each monozygotic twin pair as the index twin and calculating the association between Southeast Asia service and combat exposure with each of the five alcohol drinking measures. This yields the association of Southeast Asia service and combat exposure with alcohol drinking, unadjusted for co-twin effects. Hypothesis testing for these contingency tables uses standard Chi-square methods and the Mantel extension to test for a linear trend in proportions when appropriate.[18]

The potential confounding effect of co-twin alcohol drinking on the association between index twin alcohol drinking and military service in Southeast Asia is addressed using a logistic regression model. In this model, index twin alcohol drinking is the dependent variable while co-twin alcohol drinking and index twin military service in Southeast Asia are the independent variables. The logistic regression model, adapted for twin data, provides an estimate of the probability of an index twin having an event (e.g., being a lifetime abstainer) based on the co-twin's event history and the index twin's service in Southeast Asia or combat exposure.[19] Hypothesis testing is conducted by dividing the logistic regression coefficient for Southeast Asia service or combat exposure (entered as a five level ordinal variable) by its respective standard error; when squared this follows a Chi-square distribution with one degree of freedom.

The potential confounding effects of demographic and military service factors were examined in a logistic regression model that already included co-twin alcohol drinking. Confounding factors include: years of education at enlistment, branch of service, length of active duty military service, age at enlistment, calendar year of discharge, and score on a general aptitude test taken at enlistment. These factors were examined both singly and in combination, and did not substantially alter the results from the co-twin adjusted analysis.

Results

Table 1 displays the nearly identical distribution of the three measures of alcohol abstention and the two measures of alcohol consumption for index twins and co-twins separately.

Table 2 displays the joint frequency distribution of the alcohol drinking measures with service in Southeast Asia and combat exposure for index twins. There is no evidence of a trend in the prevalence of lifetime abstention, lifetime non-regular drinking, and current abstaining with increasing levels of combat exposure. Conversely, veterans who served in Southeast Asia tend to consume more alcohol than

TABLE 1—Distribution of Alcohol Drinking Measures for Monozygotic Twin Pairs in the Vietnam Era Twin Registry

Alcohol Drinking Measures	Index Twin (n = 2,169) %	Co-Twin (n = 2,169) %
A. Lifetime Abstainer		
Yes	6.2	6.5
No	93.8	93.5
B. Lifetime Non-Regular Drinker		
Yes	24.2	24.9
No	75.8	75.1
C. Current Abstainer		
Yes	37.3	37.2
No	62.7	62.8
D. Average Daily Ethanol Consumption		
Abstainer	37.3	37.2
Low	19.6	20.8
Moderate	24.2	23.9
Heavy	18.7	18.2
E. High Consumption		
Yes	5.5	5.4
No	94.5	94.6

TABLE 2—Distribution of Alcohol Drinking Measures by Southeast Asia (SEA) Service and Combat Exposure in Monozygotic Index Twins

	Service in Southeast Asia		Combat Exposure				
	No	Yes	Non SEA	SEA No Combat	SEA Low Combat	SEA Med Combat	SEA High Combat
	(n = 1,340)	(n = 829)	(n = 1,340)	(n = 196)	(n = 222)	(n = 245)	(n = 166)
Alcohol Drinking Measures	%	%	%	%	%	%	%
A. Lifetime Abstainer							
Yes	6.6	5.6	6.6	7.6	3.2	5.3	6.6
No	93.4	94.4	93.4	92.4	96.8	94.7	93.4
	(X^2 = 0.916, p = 0.339)			(X^2_{trend} = 0.835 with 1 d.f., p = 0.361)			
B. Lifetime Non-Regular Drinker							
Yes	24.6	23.4	24.6	28.6	22.1	20.0	24.1
No	75.4	76.6	75.4	71.4	77.9	80.0	75.9
	(X^2 = 0.419, p = 0.517)			(X^2_{trend} = 1.352 with 1 d.f., p = 0.245)			
C. Non-Current Drinker							
Yes	38.4	35.5	38.4	39.8	33.3	33.1	36.8
No	61.6	64.5	61.6	60.2	66.7	66.9	63.2
	(X^2 = 1.929, p = 0.165)			(X^2_{trend} = 2.395 with 1 d.f., p = 0.122)			
D. Average Daily Ethanol Consumption							
Abstainer	38.4	35.5	38.4	39.8	33.3	33.1	36.8
Light	19.6	20.1	19.6	21.9	22.5	16.7	19.9
Moderate	24.9	23.0	24.9	23.0	23.4	25.7	18.7
Heavy	17.0	21.4	17.0	15.3	20.7	24.5	24.7
	(X^2 = 3.738, p = 0.053)			(X^2_{trend} = 6.625 with 1 d.f., p = 0.010)			
E. High Consumption							
Yes	4.6	7.0	4.6	5.6	5.0	8.6	9.0
No	95.4	93.0	95.4	94.4	95.0	91.4	91.0
	(X^2 = 5.898, p = 0.015)			(X^2_{trend} = 9.320 with 1 d.f., p = 0.002)			

veterans who did not serve in Southeast Asia (p = 0.053). The prevalence of light and moderate consumption is similar regardless of where a veteran served, but there are 4.4 percent more Southeast Asia veterans than non-Southeast Asia veterans who are heavy alcohol consumers. The extent of combat exposure is directly associated with increased average daily ethanol consumption. Non-Southeast Asia veterans and Southeast Asia veterans who experienced no combat consume about the same amount of alcohol, but veterans who have been exposed to either medium or high combat are found to consume more alcohol. When high consumption is examined, a 2.4 percent excess of high consumers is found in Southeast Asia veterans compared to non-Southeast Asia veterans. Further, a nearly monotonic gradient is observed in the prevalence of high alcohol consumption with combat exposure.

The association of the five alcohol drinking measures with Southeast Asia service and combat exposure, after adjustment for co-twin effects, is displayed in Table 3. None of the three abstention measures as well as average daily ethanol consumption are associated with Southeast Asia service or combat exposure. However, of all index twins who served in Southeast Asia, 5.6 percent are classified as high consumers, compared to 3.9 percent of index twins who did not serve in Southeast Asia (p = 0.048). Similarly, the ordinal index of combat exposure remains positively correlated with the co-twin adjusted prevalence of index twins who are classified as high consumers (p = 0.034).

Discussion

Our results demonstrate a complex pattern of association between military service experience and alcohol drinking patterns. The abstention measures are not associated with military service experience in either the unadjusted or adjusted analysis. In the unadjusted analysis there is a strong relationship of both measures of alcohol consumption with Southeast Asia service and combat exposure. In the cotwin adjusted analysis, average daily ethanol consumption is no longer significantly related to service in Southeast Asia and combat exposure; the association of high consumption with Southeast Asia service and combat is substantially diminished.

What is the likelihood that our findings are due to bias? The principal biases that are likely to be operating in this study relate to uncontrolled confounding, non-response, and flaws in measurement due to the self-reporting of military service experience and alcohol drinking. The principal uncontrolled confounding factor in this study is the pattern of alcohol drinking prior to entering military service. We chose not to collect measures of pre-military service alcohol drinking because of the potential of biased recall. It seems highly probable that the recollection of alcohol drinking patterns 15 to 20 years ago might be influenced by whether a veteran served in Southeast Asia.

Non-response bias might distort our results if veterans who responded are systematically different from veterans who did not respond *and* this difference is related to both alcohol drinking and military service in Southeast Asia. To gain further insight into the possibility of non-response bias we compared the association of Southeast Asia service with demographic and military service confounding factors (years of education at enlistment, branch of service, length of active duty military service, age at enlistment, calendar year of discharge, and score of the general aptitude test taken at enlistment), in responders and nonresponders. This analysis revealed that the potential for bias exists for education at enlistment and age at enlistment. In particular, education at enlistment was directly related to the probability of Southeast Asia service in responders and inversely related to the

TABLE 3—Distribution of Alcohol Drinking Measures by Southeast Asia (SEA) Service and Combat Exposure in Monozygotic Index Twins Adjusted for Co-Twin Alcohol Drinking Measures

	Service in Southeast Asia		Combat Exposure				
	No	Yes	Non SEA	SEA No Combat	SEA Low Combat	SEA Med Combat	SEA High Combat
Alcohol Drinking Measures	%	%	%	%	%	%	%
A. Lifetime Abstainer							
Yes	4.1	3.9	4.1	4.1	4.0	3.9	3.9
No	95.9	96.1	95.9	95.9	96.0	96.1	96.1
	($X^2 = 0.07$, $p = 0.789$)			($X^2 = 0.05$ with 1 d.f., $p = 0.815$)			
B. Lifetime Non-Regular Drinker							
Yes	21.5	21.6	21.8	21.5	21.2	20.9	20.6
No	78.5	78.4	78.2	78.5	78.8	79.1	79.4
	($X^2 = 0.01$, $p = 0.934$)			($X^2 = 0.200$ with 1 d.f., $p = 0.654$)			
C. Non-Current Drinker							
Yes	36.8	34.9	36.8	36.1	35.3	34.5	33.7
No	63.2	65.1	63.2	63.9	64.7	65.5	66.3
	($X^2 = 0.72$, $p = 0.396$)			($X^2 = 0.94$ with 1 d.f., $p = 0.333$)			
D. Average Daily Ethanol Consumption							
Abstainer	35.0	33.1	35.2	34.2	33.2	32.2	31.2
Light	23.8	23.6	23.8	23.7	23.6	23.5	23.4
Moderate	26.7	27.7	26.6	27.2	27.7	28.2	28.7
Heavy	14.5	15.6	14.4	14.9	15.5	16.1	16.7
	($X^2 = 1.07$, $p = 0.301$)			($X^2 = 2.29$ with 1 d.f., $p = 0.131$)			
E. High Consumption							
Yes	3.9	5.6	4.0	4.5	5.2	5.9	6.7
No	96.1	94.4	96.0	95.5	94.8	94.1	93.3
	($X^2 = 3.89$, $p = 0.048$)			($X^2 = 4.52$ with 1 d.f., $p = 0.034$)			

probability of Southeast Asia service in non-responders. Likewise, there was no difference in Southeast Asia service by age at enlistment in responders, but in non-responders there was an increased probability of serving in Southeast Asia for veterans who enlisted after the age of 19.

Bias due to self-reported military service experiences is unlikely to provide an explanation for our results. The measurement of Southeast Asia service is in accord with Southeast Asia service as determined from the military record (kappa = 0.86). We have previously demonstrated that the index of combat exposure is reliable and valid.[15] Likewise, the five alcohol drinking measures used in this study demonstrated acceptable reliabilities.

It is difficult to relate our findings to previous research on alcohol drinking among Vietnam era veterans[7,8,11,20,21] because of the major methodological differences among studies. The definition of "exposed" and "unexposed" cohorts is highly variable across studies. Some compare Vietnam era veterans with non-veterans without any measure of actual service in Southeast Asia[22] while others compare combat exposed and non-exposed Southeast Asia veterans without any non-veteran comparison group.[20] Each of the studies uses a different set of indicators for alcohol drinking behavior, ranging from a diagnosis of alcohol abuse and dependence,[8] to measures of usual alcohol beverage consumption.[21] The time interval between exposure and the measurement of alcohol drinking behavior also displays great variability across studies. For example, Robins'[10] initial data collection took place within two years of discharge from active duty while Kulka, *et al*,[12] collected data more than 15 years following discharge. Finally, and perhaps most importantly, there are major differences in the way each study addresses the issue of confounding. If factors related to both military service and alcohol consumption are not adequately controlled in the study design or analysis, then the observed results could be wholly or partially confounded.

Boscarino,[21] using data collected in 1977, found a positive association between service in Southeast Asia and three measures of alcohol consumption (usual daily intake, heavy drinking during the past 30 days, and binge drinking). That study is especially troublesome since it failed to exclude females from the analysis and the confounder adjustment was based solely on demographic factors. In a recent study, the Centers for Disease Control[8] found an increased prevalence of alcohol abuse and dependence in Southeast Asia and combat exposed veterans compared with non-Southeast Asia veterans. That study controlled for potential confounders by restricting the study participants to male Army veterans and by matching Southeast Asia veterans to non-Southeast Asia veterans based on period of service and age. During the analysis numerous potential confounders were examined and found to have no effect. However, the analysis failed to account for familial alcohol drinking patterns.

Card[9] used a single question on heavy alcoholic beverage intake to compare an age-matched cohort of non-veterans, veterans who did not serve in Southeast Asia, and veterans who served in Southeast Asia. Using a complex research design that involved matching on over 50 premilitary service factors, she found no difference in the prevalence of heavy drinking according to service in Southeast Asia. Roth[11] found a small but significant association between military service in Southeast Asia and a composite measure of alcohol intake and problem drinking, after multi-factor adjustment; she found no association between combat exposure and the same composite drinking measure. Yager, *et al*,[7] using a slightly different measure of alcohol drinking in this same cohort, came to a different conclusion; neither an indicator of heavy daily alcohol consumption nor a problem drinking scale was related to service in Southeast Asia. A significant association was found between combat and a measure of heavy daily consumption. These conflicting reports obtained from the same data decrease the credibility of both studies.

Based on our findings, and the inconsistent evidence from previous research, we conclude that service in Southeast Asia and exposure to combat have relatively modest long-term effects on the alcohol drinking patterns of Vietnam era veterans.

APPENDIX A

Eighteen Indicators of Combat Exposure

1. In an artillery unit which fired on the enemy
2. Flew in an aircraft (reconnaissance, or fixed wing F-14, B-52, etc.)
3. Flew helicopter attack gunships or medivacs
4. Stationed at a forward observation post
5. Tunnel rat checking enemy base camps
6. Served on river patrol or gunboat
7. Demolitions expert in the field
8. Assigned to graves and registration to retrieve dead bodies from the field
9. Served as a medic in combat
10. Received incoming fire
11. Encountered mines and booby traps
12. Received sniper or sapper fire
13. Unit patrol ambushed
14. Flew in aircraft (fixed wing or helicopters) and was shot down
15. Engaged VC and/or NVA in firefight
16. Saw Americans killed, and/or saw Vietnamese killed
17. Wounded
18. Captured by the enemy

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Violence during Pregnancy and Substance Use

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Abstract: Violent incidents were assessed as part of a prospective study of 1,243 pregnant women. Participants were predominantly poor, urban, minority group women. Seven percent (n = 92) of women reported physical or sexual violence during pregnancy. Most of the women (94 percent) knew their assailant. Victims of violence were at greater risk of having a history of depression and attempted suicide, having more current depressive symptoms, reporting less happiness about being pregnant, and receiving less emotional support from others for the current pregnancy. Comparisons of victims and non-victims showed that victims were more likely to be users of alcohol and drugs. In addition, partners of victims were more likely to use marijuana and cocaine. When possible confounders were controlled using multivariable analyses, a woman's alcohol use during pregnancy and her partner's drug use were independently associated with an increased risk of being a victim of violence during pregnancy. Results of this study highlight the importance of assessing exposure to violence during prenatal care, especially among women who are heavy users of alcohol or drugs or whose partners use these substances. (*Am J Public Health* 1990; 80:575–579.)

Introduction

A substantial proportion of women in the general population experience abuse and violence.[1-8] Two recent studies[6,7] indicate that 4 to 8 percent of pregnant women experience abuse during pregnancy and that this problem may be largely unreported to health care providers. Since the risk of abuse is higher for women in young adulthood and this time coincides with women's most active childbearing years,[9] the experience of violence and abuse during pregnancy deserves special attention.

In spite of the prevalence of violence experienced by women, and its damaging impact on women's health and mental health, few studies have attempted to identify the factors associated with violence during pregnancy. One such study found little difference between abused and non-abused pregnant women in race, age, marital status, employment or educational status.[6] Another study[7] found that, compared to non-abused women, women who experienced abuse during pregnancy were of lower socioeconomic status, higher parity, more likely to be single, and had a history of depression or other psychiatric symptoms. Abused women were also more likely to report alcohol use, and a trend in the data indicated that they may also have higher levels of illicit drug use.[7] While this preliminary finding is consistent with previous associations between alcohol and drug use and family violence in the general population,[8] methodological problems such as small numbers of abused women in the sample and reliance on self-reported drug use[10] limit firm conclusions about the role of these factors during pregnancy.

The aims of the present analysis are: to describe the prevalence and patterns of violent incidents during pregnancy; to describe the association between demographic and psychosocial characteristics and violence during pregnancy; to investigate the association between the experience of violence during pregnancy and the use of alcohol and illicit drugs by pregnant women and their partners; and to investigate the association between the experience of violence during pregnancy and newborn outcomes.

Methods

Sample Recruitment

Subjects were consecutively recruited from July 14, 1984 through June 30, 1987 in the Women's and Adolescent Prenatal Clinics of Boston City Hospital. Eligible subjects were English- or Spanish-speaking women who were willing to give informed consent for study participation. The protocol was approved by the Human Studies Committee of Boston City Hospital. Subjects were protected from the use of data for criminal prosecution by a Writ of Confidentiality obtained under Title 42 of the United States Code Section 242A.

Assessment of Participants

All participants were interviewed during the prenatal and postpartum period by a trained bilingual interviewer. To minimize attrition, participants were paid $10 for each interview. The first interview assessed violent incidents occurring in the time period from three months prior to the calculated day of conception through the day of the interview. The postpartum interview determined violent incidents from the time of the previous interview through the time immediately prior to the delivery of the baby. Violent incidents were ascertained by asking the following question, "Were you physically threatened or abused, or were you involved in any fights or beatings?" The respondents were then asked to "describe what happened," and probed to find out whether the assailant was known to the victim, the type of injury, whether they had seen a doctor, and whether they were hospitalized overnight. Victims of violence were defined as those women who experienced physical or sexual violence during their pregnancies. Verbal threats or emotional abuse were not included as an episode of violence in this analysis.

A close-end, forced-choice interview elicited sociodemographic characteristics and the timing and frequency of the use of cigarettes, alcohol, marijuana, cocaine, and other illicit psychoactive substances prior to and during pregnancy. Data on drug and alcohol use were obtained for the same time periods as were data on episodes of violence. Participants were asked to furnish urine samples at the time of each interview and were informed that their urine would be assayed for "marijuana metabolites and prescription and nonprescription drugs," but that these results would remain confidential and not become part of their clinical record. The

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use of other substances such as alchohol and other drugs was ascertained by self-report alone. Urine assays were conducted by Enzyme Mediated Immunoassay Technique (EMIT),[11] and positive results were confirmed by Mass Spectrometry/Gas Chromatography for cocaine metabolites and by High Pressure Liquid Chromatography[12] for marijuana metabolites. Use of these drugs was determined by both positive self-report or positive urine assay. A full description of the procedure for the urine assays has been published elsewhere.[13]

During the prenatal interview, respondents were also asked to report whether their partners used alcohol, marijuana or cocaine. Quantity measures of partner's drug use were also asked but proved to be difficult for the women to answer and were therefore not included in the analysis.

Life stress was measured by the Life Experiences Survey (LES) whose reliability and validity has been demonstrated.[14] For this analysis, life stress was measured by the Negative Life Events subscale of the LES, which is a summary scale of the number and negative effect of life events experienced in the past year.

Depressive symptoms were measured by the Center for Epidemiological Studies—Depression Scale (CES-D)[15] whose validity and reliability for the CES-D have been demonstrated and reported.[15–22] Past history of depression and suicide were measured by asking respondents, "In your lifetime, have you had two weeks or more during which you felt sad, blue, depressed, or when you lost all interest and pleasure in things that you usually cared about or enjoyed?" History of suicidal feelings and attempts were assessed by asking "Has there ever been a period of two weeks or more when you felt like you wanted to die?" and "Have you ever attempted suicide?"

Following delivery, trained record reviewers, without knowledge of the hypotheses of this analysis, abstracted medical records using a precoded protocol, which documented participants' reproductive and general medical health histories.

Statistical Analysis

Odds ratios and 95% confidence intervals (CI) were used to compare victims of physical or sexual violence during pregnancy to non-victims of violence on demographic, medical and psychosocial characteristics and on drug and alcohol use.[23] Least square multiple regression was used to determine the association between neonatal growth parameters and violence during pregnancy while controlling for potentially confounding variables.[24] Multiple logistic regression was used to assess the multivariable relationship between violence and drug use. Adjusted odds ratios and 95% CI were calculated from this model.[23]

Results

Sample Characteristics

Eighty-six percent of 1,932 eligible women participated in the study; 144 (8 percent) refused to participate and 124 (6 percent) who had agreed to participate left after the prenatal clinic visit prior to the interview. Non-participants were more likely than participants to be White (19 vs 9 percent), older (26 years vs 24 years), cigarette smokers (41 vs 34 percent) or third trimester registrants (40 vs 22 percent). The two groups were similar in parity, use of alcohol, marijuana, cocaine, or other drugs.

Seventy-eight percent (n = 1,303) of the women who were recruited completed the study. Of the 361 women who were lost to follow-up, most (81 percent) delivered in another hospital or had an elective abortion. Sixty (5 percent) of the 1,303 who completed the study were excluded from the analysis because of a missing medical record review or infant examination, leaving 1,243 completed cases. Completed cases (n = 1,243), when compared to those lost to follow-up or with incomplete infant data (n = 421), had similar sociodemographic characteristics and similar patterns of alcohol, marijuana, cocaine, and other drug use.

The sample of 1,243 participants with complete data reflects the low income (48 percent with household income ≤ $500 a month), predominantly ethnic minority (North American Black = 55 percent, foreign-born Black = 19 percent, Hispanic = 18 percent, White = 8 percent), single (62 percent), and primiparous (55 percent) population served at the prenatal clinic. While 16 percent of the participants were less than 18 years old, 66 percent were between 19–29 years of age.

Violence during Pregnancy

Seven percent (n = 92) of women in the study reported physical or sexual violence during pregnancy. An additional 3 percent of the study population (n = 37) reported incidents of violence in the three months prior to the pregnancy but not during their pregnancy. Less than one percent (n = 11) of women in the sample (12 percent of victims) reported incidents both three months before and during pregnancy.

Sixty percent of the victims were subjected to one incident of violence during pregnancy, 25 percent were victimized twice and 15 percent experienced three or more incidents. More incidents occurred in the first trimester (55 percent) than in the second (40 percent) or third trimesters (25 percent). Most of the women (94 percent) knew their assailant. Thirty-six percent of the victims saw a doctor for at least one of the violent incidents during pregnancy and 10 percent were hospitalized overnight as a result of one of these incidents.

Comparison of Victims and Non-Victims

Compared to non-victims (Table 1), victims had higher relative odds of identifying themselves as White (7 vs 15 percent), born in the US (65 vs 76 percent), and single (62 vs 75 percent). They were also more likely to be on Medicaid (60 vs 76 percent), to have had a history of sexually transmitted diseases (17 vs 30 percent) and to have had an elective

TABLE 1—Relative Odds of Selected Demographic and Medical Differences of Victims of Violence Compared to Non-victims (n = 1,243)

Variable/Comparison	Relative Odds	95% Confidence Interval
Ethnicity		
North American Black/Other Black	1.56	0.81, 3.00
North American Black/Hispanic	1.54	0.78, 3.03
North American Black/White	0.48	0.25, 0.92
North American Black/Other	0.75	0.37, 0.54
Marital Status		
Living with Partner/Married	1.86	0.80, 4.34
Single/Married	2.52	1.28, 4.97
Born in the US/Born outside	1.70	1.04, 2.79
Medicaid Recipient/Non-recipient	2.12	1.23, 3.66
History of Sexually Transmitted Disease*/None	2.13	1.33, 3.41
Prior Elective Abortion/None	1.68	1.01, 2.57

*Syphilis, gonorrhea, herpes, pelvic inflammatory disease, chlamydia abortion (38 vs 50 percent). The two groups had similar levels of educational attainment (42 percent of victims had 12 years of school vs 39 percent of non-victims), mean ages ($x = 24$ for each group), employment status (24 percent not employed in each group) and parity (53 percent of victims were primiparas vs 52 percent of non-victims).

Victims of violence differed from non-victims on a number of psychosocial measures (see Table 2). Compared to non-victims, victims were at greater risk of having reported a history of one or more depressive episodes (33 vs 56 percent) and to have attempted suicide (5 vs 17 percent), reporting unhappy feelings about their current pregnancy (20 vs 33 percent), perceiving their partners (11 vs 22 percent) and families (14 vs 27 percent) as unhappy about their pregnancy, and feeling a lack of support during their pregnancy (13 vs 36 percent). Victims reported more depressive symptoms (mean difference = 5.6, 95% CI = 3.35, 7.86) and greater numbers of negative life events in the past year (mean difference = 13.16, 95% CI = 10.01, 16.31).

These demographic and psychosocial characteristics did not differ between women who experienced one compared to two or more incidents of abuse during pregnancy (data available on request to authors).

Drug Use among Victims and Non-victims

Women were divided into three categories (nonusers, light users, and heavy users) based on their frequency of use and the number and type of different drugs used during pregnancy. Since we have shown that among marijuana and cocaine users, a positive urine assay is associated with more frequent drug use,[13] we designated women who had a positive urine for either drug and those who self-reported using opiates and/or marijuana or cocaine at least weekly during pregnancy as heavy users. Users of any illicit drug who did not meet the heavy use criteria were categorized as light users.

Victims of violence during pregnancy were at greater risk than non-victims of being heavy users of alcohol (OR = 2.43, 95% CI = 1.71, 3.46) and illicit drugs (OR = 2.68, 95% CI = 1.72, 4.17). A comparison of frequency of use by victims and non-victims showed that victims were heavier substance users in all categories of use than non-victims (see Table 3).

TABLE 2—Relative Odds of Selected Psychosocial Differences of Victims of Violence Compared to Non-victims (n = 1,243)

Variable/Comparison	Relative Odds	95% Confidence Interval
History of Depression		
Yes/No	2.60	1.69, 4.00
History of Suicide Attempts		
Yes/No	3.85	2.08, 7.14
Unhappy Feelings about Pregnancy		
Yes/No	1.91	1.21, 3.02
Partner's Unhappy Feelings about the Pregnancy		
Yes/No	2.40	1.37, 4.19
Family's Unhappy Feeling about the Pregnancy		
Yes/No	2.62	2.29, 5.73
Little or No Emotional Support During the Pregnancy		
Yes/No	2.30	1.37, 3.86

TABLE 3—Relative Odds of Frequency of Psychoactive Substance Use of Victims of Violence Compared to Non-victims (n = 1,243)

Variable/Comparison	Relative Odds	95% Confidence Interval
Maximum Average Daily Alcohol Use		
< 1 Drink/None	2.06	1.24, 3.44
1–2 Drinks/None	5.34	2.30, 12.40
2+ Drinks/None	5.15	2.14, 12.41
Frequency of Marijuana Use[a]		
< 1× a Month/None	1.98	0.81, 4.81
Monthly/None	2.13	0.93, 4.91
Weekly/None	2.56	1.53, 4.29
Frequency of cocaine use[1]		
< 1× a Month/None	1.67	0.70, 4.03
Monthly/None	2.69	0.90, 8.02
Weekly/None	2.91	1.56, 5.44

[a] Based on self-reported data only.

Compared to non-victims, women who were victims of violence had greater odds of having a male partner who was a marijuana (OR = 2.27, 95% CI = 1.46, 3.53) and/or cocaine (OR = 2.35, 95% CI = 1.81, 4.48) user. Further, odds ratios demonstrate a two-fold increase in the use of two or more drugs (OR = 3.2, 95% CI = 1.92, 5.48) by partners of victims when compared to partners of non-victims. Alcohol use by partners was not an increased risk for victims (OR = 1.03, 95% CI = 0.63, 1.69).

Alcohol and drug use did not differ between women who experienced one versus two or more incidents of abuse during pregnancy (data available on request to authors).

A multiple logistic regression (Table 4) shows that the risk of being a victim of violence was associated with a woman's alcohol use during pregnancy and drug use by her partner even when controlling for race, age, marital status, education, and history of violence in the three months prior to pregnancy. While heavy drug use during pregnancy (use of opiates or weekly use of marijuana or cocaine or positive urine assay) was associated with a 39 percent increase in the odds of being a victim of violence, chance cannot be ruled out as the explanation for this finding. Violent incidents in the three months prior to pregnancy have strong predictors of violence during pregnancy.

Birth Outcomes

Multivariable regression analyses were conducted in

TABLE 4—Results of Multiple Logistic Regression Analysis on Experience of Violence During Pregnancy (n = 1,153)

Variables	Relative Odds	95% Confidence Interval
Average daily alcohol use		
(2 drinks/day vs none)	1.87	(1.24, 2.80)
Average daily alcohol use		
1 drink/day vs none)	1.37	(1.12, 1.67)
Heavy illicit drug use	1.39	(0.79, 2.46)
Light illicit drug use	0.87	(0.36, 2.12)
Partner's alcohol use	1.41	(0.82, 2.43)
Partner's illicit drug use		
(2 or more vs none)	2.26	(1.19, 4.30)
Partner's illicit drug use		
(1 vs none)	1.51	(1.09, 2.07)
White vs US Black	1.95	(0.97, 3.93)
Foreign-born Black vs US Black	0.86	(0.41, 1.77)
Hispanic vs US Black	1.02	(0.49, 2.11)
Other race vs US Black	0.72	(0.20, 2.59)
Age (30 years vs 18 years)	0.90	(0.51, 1.60)
Marital status (married vs single)	0.60	(0.29, 1.23)
High school graduate vs not	1.12	(0.67, 1.85)
Violent incidents—3 months prior to pregnancy	5.87	(2.73,12.60)

order to investigate the relationship between violence during pregnancy and birthweight, infant length, head circumference, and gestational age. Each regression model included ethnicity, age, pre-pregnancy weight, weight gain during pregnancy, parity, number of prenatal visits, history of sexually transmitted diseases, marijuana use, cocaine use, alcohol use, opiate use, and cigarette use. Once confounders were controlled, the association of violence to birth outcomes, while generally in the expected direction, was weak: birthweight: −19 g, (95% CI = −115, 78); length: 0.15 cm, (95% CI = −0.68, 0.38); head circumference = −0.14 cm, (95% CI = −0.48, 0.19); gestational age: 0.13 weeks, 95% CI = −0.29 to 0.55).

Discussion

The findings of this study support previous reports on the prevalence of violence during pregnancy.[6,7] Of greater importance are the results which provide evidence of a strong relation between violence during pregnancy and the use of alcohol by the pregnant woman and the use of illicit drugs by her male partner.

A number of previous studies have shown that abused women are at greater risk of using heavier quantities of alcohol and prescription and non-prescription drugs than women who are not abused.[4,7] While the cause and effect relationship between use of alcohol and experience of violence is unclear, there are indications that women who are abused may self-medicate with alcohol, illicit drugs, and prescription medication in order to cope with the violence.[4,5,25] For example, in a review of medical records, Stark, *et al*,[4] determined that increased alcohol and drug use followed the first incident of abuse.

The relation between the male partner's drug use and violence during pregnancy has not received as much attention. In general, among men who batter their female partners, substance use has been found to frequently accompany battering.[7,26–28] One author[29] has suggested that drinking is used by men who batter to disavow their violent behavior and disclaim responsibility for their actions. Findings from the present study indicate that multiple drug use among male partners is independently associated with over a two-fold increase in women's experience of violence during pregnancy. Since this study did not directly investigate whether the assailant was the male partner, it is not possible to identify the perpetrator of violence as the male partner. Anecdotal evidence from the interviews suggests that a large proportion of the perpetrators were male partners, however. Further research is needed to determine whether the use of illicit drugs actually promotes violent behavior, is employed as a culturally sanctioned justification for engaging in violence without taking responsibility for it, or is associated through some third causal factor to violence.

The results of this study indicate that once sociodemographic factors and drug use are taken into account, the experience of violence during pregnancy is weakly, if at all, associated with newborn size or gestational age. It is possible that infrequent but severe injuries are associated with adverse pregnancy outcomes. However, the small number of women reporting such injuries did not allow us to assess this possibility.

A number of important limitations should be considered when interpreting the results of this study. First, the subjects were primarily poor, Black, and Hispanic women living in the inner-city. The prevalence of violence and drug use in other populations may be different. Second, our assessment of abuse was limited because it did not evaluate the full psychological impact of an abusive relationship that typically involves a climate of threat and fear. Third, the assessment of violence through limited questions about current experiences of violence does not allow for an understanding of the causal ordering of the relationship between violence, drug use, and depression. Finally, validity of the assessment of the male partner's drug use may be reduced since it was conducted through the female participant, rather than through direct self-report and/or urine assay from the male partner.

With these limitations in mind, prenatal health care providers should assess women's exposure to violence prior to and during pregnancy. In addition, assessment of psychoactive drug use among women and their partners is critical, since such behaviors may be employed as a marker for risk for violence during pregnancy. Further, the possibility of continued violence against a woman after delivery and the potential extension of this violence to the child have important public health implications. Since violence and drug use reflect a wider set of environmental conditions and psychosocial factors that create a context that is detrimental to a mother's and infant's physical and mental health, prevention and intervention efforts will need to be set within a framework of services to address these needs. Training of health care professionals about the problems of drug abuse and family violence is a critical and necessary first step in early identification and intervention efforts.

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NCI Announces Research Opportunities in Cancer Prevention

The Division of Cancer Prevention and Control (DCPC) of the National Cancer Institute has announced it is accepting applications for the next Cancer Prevention Fellowship Program. Funding permitting, as many as 10 fellows will be accepted for up to three years of training, beginning July 1, 1991. The *deadline* for applications is September 1, 1990.

The purpose of this program is to attract individuals from a multiplicity of health science disciplines into the field of cancer prevention and control. The program provides for:

- Participation in the DCPC Cancer Prevention and Control Academic Course;
- Working at NCI directly with individual preceptors on cancer prevention and control projects;
- Field assignments in cancer prevention and control programs at other institutions.

Benefits include selected relocation and travel expenses, paid federal holidays, and participatory health insurance. Those eligible for the program include those with MD or DDS degree from a US, territorial, or Canadian medical school. Foreign medical graduates must have current ECFMG/FMGEMS certification and appropriate experience (e.g., one year residency) in a training program approved by the Accreditation Council for Graduate Medical Education. Also eligible are those with PhD, DrPH, or other doctoral degree in a related discipline (epidemiology, biostatistics, and the biomedical, nutritional, public health or behavioral sciences). Foreign education must be comparable to that received in accredited US, territorial, or Canadian institutions. Applicants must be US citizens or resident aliens eligible for citizenship within four years.

For more details and application catalog, send a postcard or letter with your name and home address to: Douglas L. Weed, MD, MPH, PhD, Director, CPFP, DCPC, National Cancer Institute, Executive Plaza South, T-41, Bethesda, MD 20892. Further inquiries should be directed to Barbara Redding at NCI; (301) 496-8640 or -8641.

Risk Factors for Depressive Symptomatology in a Drug Using Population

JOHN C. BUCKNER, PHD, AND WALLACE MANDELL, PHD, MPH

Abstract: This study employs a prospective design to examine possible personality, drug use, stressful life event, and social support-related variables associated with the onset of a depressive episode in a cohort of psychoactive drug using young adults. Two waves of data, collected one year apart, were available on 942 individuals. Cases (n = 62) were free of depressive symptoms at time 1 but reported significant symptomatology at time 2 as measured by the depression subscale of the Brief Symptom Inventory. Controls (n = 490) were those free of depressive symptoms at both time points. In multivariate analyses, users of the central nervous system depressant methaqualone had a nearly four-fold elevated risk for depressed mood as compared to nonusers. Additional risk factors significant after multivariate adjustment included lower self-esteem at time 1 and negative life events. These results highlight the multifactorial nature of depressive symptomatology. (*Am J Public Health* 1990; 80:580–585.)

Introduction

Using predominantly case-control comparisons on general population samples, the association between negative life events,[1–9] lack of social support,[10,11,4,5,8,12] and self-esteem[13–18] with depressive symptoms and disorder has been well documented. Analogous prospective studies are rare.[17] While research has documented an association between alcohol use and depression,[19–21] much less is known about the potential effects of other drugs or about the relative and possibly interactive effects of life events and drug use on depression symptomatology. One obstacle to such studies is the low base rate of negative life events and/or heavy substance use in most samples making it unfeasible to examine their relationships. By focusing on a population of young adults who frequently use a variety of psychoactive substances, we estimated the adjusted relative risk of various drug use, personality, psychosocial, and negative life event variables for developing an episode of depressed mood.

Methods

We utilized a case-control methodology nested within a prospective study[22,23] to investigate risk factors for depressed mood using data from a study on the consequences of arrest for marijuana possession, a short-term longitudinal investigation of drug using young adults.*

Study Population

In 1981–82, 1,134 adults living in the State of Maryland were interviewed and 942 were relocated and agreed to a reinterview one year later. A primary aim was to evaluate the effects of having been arrested and tried for marijuana possession on subsequent drug use and mental health. An application for a certificate of confidentiality, which protects the identity of research study subjects, was approved by the US Secretary of Health and Human Services. Written consent was obtained of all research participants and each individual was informed as to the nature of the study and the certificate which protects the information they provide from seizure by police or use in any civil or criminal proceeding.

The panel of 942 individuals which forms the basis of the present study was comprised, in part, of a random sample of 621 individuals, stratified by month of arrest, from a roster of all adults who had been tried in 1981 for marijuana possession within the district courts of Maryland. An additional 321 non-arrested individuals, nominated by their arrested counterparts as being friends or acquaintances, were selected as a neighborhood-based comparison group. This nomination procedure was done to ensure comparability between the two groups on demographic characteristics and drug use patterns.

Operationalization of Variables

The presence of depressive symptomatology in the week prior to interview was assessed using the depression subscale of the Brief Symptom Inventory (BSI)—the short form of the Symptom Checklist-90 (SCL-90).[24,25] Internal consistency analysis of the depression sub-scale revealed coefficient alpha = .78 at the time of first interview (Time 1) and alpha = .80 at the second interview (Time 2). The depression subscale, consisting of six items, has been found to correlate (.95) with its corresponding longer version on the SCL-90.[25] There is evidence in support of the temporal stability and the convergent and discriminant validity of the BSI depression subscale.[26]

Individuals were classified as a case if they did not report depressive symptoms at Time 1 but reported symptoms at Time 2 that placed them in the 90th centile or above on the non-patient norms for the BSI depression subscale.[26] This cut-point was selected to serve as a strong contrast between cases and controls while still retaining a sufficient sample size for purposes of statistical power. While this depression measure does not identify a case of dysthymia or major depressive disorder according to DSM-III-R[27] criteria, an individual defined as a case in this study would present clinically with at least moderate symptoms of depression. All individuals who reported no or minimal symptoms of depression at both time points were selected to form a control group. Using these criteria a total of 62 cases and 490 controls were selected. Individuals excluded from analyses were those who reported mild-to-severe depressive symptoms at Time 1. Inspection of the depression scores at both time points revealed that those selected as cases went, on average, from the 50th percentile at Time 1 to the 95th percentile at Time 2. In contrast, those selected as controls were at about the 45th percentile at Time 1 and remained there at Time 2.

The wording of the BSI allowed for a determination of the current level of depressed mood experienced by a study participant, although, it could not help determine depressed

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*Mandell W. Cuthbert B, Dykacz J, Brown CH, Seidler A, Lunday J: Consequences of arrest for marijuana possession. Unpublished manuscript, Johns Hopkins University, Baltimore, MD, 1988.

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mood in the months prior to each administration. While an imperfect and somewhat less specific measure for determining case status than use of an instrument such as the Diagnostic Interview Schedule,[28] it is unlikely that a bias was produced by this manner of case-control selection. Because the BSI was administered at Time 1, some control for prior depressed mood was achieved; more than is the case in many other case-control comparisons.

Participants were asked about frequency of use of a wide variety of licit and illicit drugs in the two months prior to each interview. Alcohol and marijuana were used with enough regularity to justify coding on an interval level (3 = daily; 2 = two–three times a week; 1 = one time a week or less; 0 = no use). Tobacco use was measured in packs of cigarettes smoked per day. The vast majority of use of such substances as amphetamines, barbiturates, methaqualone (Quaalude, Sopor), cocaine, heroin, hallucinogens, and anxiolytics (e.g., Valium, Librium) occurred on a weekly or less frequent basis. Therefore, these drug variables were dichotomously coded with 1 representing any reported use of the drug during the two months prior to interview and 0 indicating no use.

The occurrence of any of five negative stressful life events (arrest, conviction, hospitalization, traffic accident, fired from job) in the 12 months between interviews was assessed and a summative count score was created. Prior research has shown that weighting of events before creation of an overall score does not increase prediction of the dependent variable.[29–32] These negative life events are typical of those found on life event inventories[33] and represent life stressors that occur with relative frequency in this population.

Separate questions tapping perceived social support, or social embededness,[34] with family members and with peer group were asked of all respondents at both interviews: "Do you feel a sense of solidarity, closeness, belonging to your family?" Four dichotomous variables were created from the Time 1 and Time 2 questions representing perceived social embededness with family or friends or lack of support/embededness. The presence of a spouse or living together with a partner was also determined and dichotomously coded as another important indicant of social support.[32,35]

Finally, the Rosenberg Self-Esteem Scale,[36] a widely used 10-item measure of self-esteem, was administered at both interviews. Extensive research has demonstrated the predictive and construct validity of this instrument.[37] With the present sample, the instrument displayed good internal consistency (Coefficient alpha = .75) and the one year test-retest stability was r(938) = .49.

Univariate logistic regression models that estimate the crude relative risk ratio for the development of an episode of depression were first calculated for variables from each of the demographic, drug use, life events, social support, and personality domains. Using multivariate logistic regression techniques, variables predictive of case status in univariate analyses were simultaneously evaluated to estimate adjusted relative risks. The presence of multiplicative interaction was examined for pairwise combinations of variables that remained significant in the multivariate analysis.

Results

Descriptive statistics for the 62 depressed mood cases and 490 controls and crude relative risk estimates based on univariable logistic regression analyses are presented in Table 1 for categorical variables and Table 2 for variables measured on a continuous scale. Cases were fairly similar to controls with respect to age, race, sex, marital status, being employed or in school, recent history of marijuana arrest, and socioeconomic status (not shown in tables). Univariate relative risk (RR) estimates for the development of depressive symptomatology were found for any use during the two months prior to the second interview for methaqualone, anxiolytics, and less strongly for heroin, tobacco, and hallucinogens.

TABLE 1—Estimated Relative Risk for an Episode of Depressed Mood Based on Univariate Logistic Regression Analysis with Categorical Variables

Variables	Cases No. (%) (N = 62)	Controls No. (%) (N = 490)	Referent Category	Relative Risk Estimate (unadjusted)	(95%CI)	p value
Sociodemographic						
Male	52 (84)	382 (78)	female	1.47	(.72, 2.99)	.29
White	24 (39)	166 (34)	black	1.23	(.72, 2.12)	.45
Employed or in school	40 (65)	322 (66)	no	.94	(.54, 1.64)	.84
Recent marijuana arrest	45 (73)	321 (66)	no	1.39	(.77, 2.51)	.27
Drug Use						
Amphetamines (Time 1)	7 (11)	49 (10)	no use	1.15	(.49, 2.65)	.75
Amphetamines (Time 2)	9 (15)	52 (11)	no use	1.43	(.67, 3.06)	.36
Barbiturates (Time 1)	5 (8)	32 (7)	no use	1.26	(.47, 3.35)	.65
Barbiturates (Time 2)	6 (10)	28 (6)	no use	1.77	(.70, 4.45)	.3
Methaqualone (Time 1)	4 (6)	31 (6)	no use	1.02	(.35, 2.99)	.97
Methaqualone (Time 2)	10 (16)	23 (5)	no use	3.90	(1.76, 8.65)	.0008
Anxiolytics (Time 1)	3 (5)	29 (6)	no use	.81	(.24, 2.74)	.73
Anxiolytics (Time 2)	8 (14)	26 (5)	no use	2.64	(1.14, 6.12)	.02
Cocaine (Time 1)	16 (26)	96 (20)	no use	1.43	(.77, 2.63)	.25
Cocaine (Time 2)	14 (23)	99 (20)	no use	1.15	(.61, 2.17)	.66
Hallucinogens (Time 1)	6 (10)	26 (5)	no use	1.91	(.75, 4.84)	.17
Hallucinogens (Time 2)	7 (11)	25 (5)	no use	2.37	(.98, 5.73)	.06
Heroin (Time 1)	4 (6)	16 (3)	no use	2.04	(.66, 6.32)	.21
Heroin (Time 2)	7 (11)	24 (5)	no use	2.47	(1.02, 6.00)	.046
Social Support (Time 2)						
Family Embededness (yes)	53 (85)	453 (93)	no	.41	(.19, .91)	.03
Peer Group Embededness (yes)	42 (68)	400 (82)	no	.47	(.27, .85)	.01
Married or Living Together	10 (16)	120 (24)	no	.59	(.29, 1.20)	.15

TABLE 2—Estimated Relative Risk for an Episode of Depressed Mood Based on Univariate Logistic Regression Analyses with Continuous-scale Variables

Variables	Cases (N = 62) Mean (S.D.)	Controls (N = 490) Mean (S.D.)	Relative Risk Estimate (unadjusted)	(95%CI)	p value
Sociodemographic					
Age	24.50 (6.21)	24.92 (6.88)	.99	(.96, 1.04)	.92
Drug Use					
Alcohol (Time 1)*	1.51 (1.01)	1.37 (.96)	1.16	(.88, 1.53)	.28
Alcohol (Time 2)*	1.62 (1.02)	1.43 (.97)	1.22	(.93, 1.61)	.15
Marijuana (Time 1)*	1.85 (1.10)	1.77 (1.12)	1.04	(.84, 1.30)	.72
Marijuana (Time 2)*	1.81 (1.25)	1.56 (1.15)	1.20	(.95, 1.51)	.12
Tobacco (Time 2)**	3.89 (2.07)	3.35 (1.86)	1.17	(1.01, 1.35)	.04
Personality (coded 1–4)					
Self-esteem (Time 1)	3.24 (.49)	3.45 (.40)	.32	(.17, .59)	.0003
Self-esteem (Time 2)	3.08 (.45)	3.46 (.42)	.14	(.07, .26)	<.0001
Life Events: (count 0–4)					
Negative Events	.79 (.91)	.44 (.72)	1.67	(1.24, 2.26)	.0009

*These frequency codes range from 0–3: (0 = no use; 1 = 1 time a week or less; 2 = 2–3 times a week; 3 = daily use).

**These frequency codes range from 1–7: (1 = no use; 2 = >1 cigarette a day, 3 = 1–5 cigarettes a day; 4 = 10 cigarettes a day; 5 = 20 cigarettes a day; 6 = 30 cigarettes a day; 7 = 40 cigarettes a day).

TABLE 3—Estimated Relative Risk for an Episode of Depressed Mood Based on Multivariate Logistic Regression Model with Drug Variables Only

Variable	Referent Category	Estimated Relative Risk	(95%CI)	p value
Drug Use (Time 2)				
Amphetamines	no use	.81	(.29, 2.30)	.69
Barbiturates	no use	.54	(.14, 7.08)	.37
Methaqualone	no use	5.27	(1.86, 14.96)	.002
Anxiolytics	no use	2.07	(.75, 5.71)	.16
Cocaine	no use	.53	(.24, 1.19)	.13
Hallucinogens	no use	2.14	(.77, 6.00)	.15
Heroin	no use	3.32	(1.18, 9.30)	.02
Alcohol frequency	NA*	1.06	(.79, 1.43)	.67
Marijuana frequency	NA*	1.07	(.82, 1.40)	.59
Tobacco frequency	NA**	1.08	(.93, 1.25)	.34

*Not applicable. These frequency codes range from 0–3: (0 = no use; 1 = 1 time a week or less; 2 = 2–3 times a week; 3 = daily use).

**Not applicable. These frequency codes range from 1–7: (1 = no use; 2 = >1 cigarette a day; 3 = 1–5 cigarettes a day; 4 = 10 cigarettes a day; 5 = 20 cigarettes a day; 6 = 30 cigarettes a day; 7 = 40 cigarettes a day).

To statistically control for multiple drug use, a multivariate logistic regression model containing main effect terms for all Time 2 drug variables (see Table 3) was created. The relative risk estimates show a clear excess risk for those reporting any use of methaqualone in the prior two months (RR = 5.27; 95% CI = 1.86, 14.96) as well as those reporting any use of heroin (RR = 3.32; 95% CI = 1.18, 9.30). Two-way interaction terms were then evaluated for methaqualone and heroin with combinations of the other drug variables. None were found to be statistically significant at ($p < .05$) although the interaction term for methaqualone and barbiturate use was significant at ($p < .10$) and was further evaluated in subsequent analyses.

Crude relative risk estimates for the life event, social support, and self-esteem measures are apparent in Tables 1 and 2. Higher levels of self-esteem at Time 1 appear to act as a protective factor with the risk of depressed mood dropping by an estimated .32 (95% CI = .17, .59) with each unit increase in self-esteem. Conversely, individuals with lower levels of self-esteem appear more vulnerable to the development of an episode of depressed mood. A strong cross-sectional relationship between Time 2 self-esteem and depression was also present. Negative life events happening in the year before follow-up served to elevate the risk for depression with each occurrence of such an event (arrest, conviction, fired from job, or hospitalization) increasing the risk for developing depressed mood 1.67-fold (95% CI = 1.24, 2.26). Risk for depressed mood was less if the subjects reported that they felt embedded within their family (RR = .41; 95% CI = .19, .91) or within a peer group (RR = .47; 95% CI = .27, .85). Being married or living with someone was not associated with case-status.

In a multivariate logistic regression model (see Table 4) containing main effect terms for life events, social support and self-esteem, the relative risk estimate for family as well as for peer group embeddedness were sharply reduced suggesting that the univariate relationships were due to associations with other variables. Self-esteem at the earlier period and the occurrence of negative events during the year remained strong predictors of case status. All possible two-way interactions among statistically significant variables in this model were tested but no interaction was detected.

We then tested a multivariable function (table not shown) that combined the drug use and psychosocial varia-

TABLE 4—Estimated Relative Risk for an Episode of Depressed Mood based on Multivariate Logistic Regression Model with Social Support, Self-esteem, and Negative Event Variables Only

Variable	Referent Category	Estimated Relative Risk	(95%CI)	p value
Family embededness (yes)	no	.50	(.21, 1.16)	.11
Peer group embededness (yes)	no	.58	(.35, 1.22)	.18
Self-esteem (Time 1)	NA*	.39	(.21, .75)	.005
Negative events	NA**	1.60	(1.15, 2.15)	.005

*Not applicable. This variable is measured on a continuous scale.
**Not applicable. This variable ranges from 0–4 and was not categorized.

bles that had retained p values of .05 or less in the prior multivariate models as well as the interaction term for methaqualone and barbiturate use. The prior association between heroin use and depressed mood no longer held in this model nor did the methaqualone-barbiturate interaction term. Negative events, self-esteem, and methaqualone use maintained their independent effects. A final model was tested that retained these three main effect terms and produced the following estimates: negative events (RR = 1.45; 95% CI = 1.06, 2.00); self-esteem (RR = .32; 95% CI = .17, .59); and methaqualone use (RR = 3.70; 95% CI = 1.59, 8.63). Interaction was not detected among pairwise combination of these three variables.

Discussion

Prior research evidence has suggested associations between self-esteem, the occurrence of negative life events, and indices of social support with depressed symptomatology. With the exception of social support, which only evidenced a trend in the predicted direction in multivariate analysis, this was corroborated in our findings; and, importantly, our results indicate that, when controlling for each other, these variables remain independent predictors for who will develop a new episode of depression. It is conceivable that a more refined measure of social support than the ones used in this study might have detected a support-depression association in multivariate, not just univariate, analyses. The measure of life events was limited by the available data to a count of five negative occurrences (arrest, conviction, hospitalization, traffic accident, fired from job). A more exhaustive index of life events would likely have produced an even stronger association.

This study has documented that a similar pattern of risk factor relationships for depressed affect exists among a sample of psychoactive drug users as has typically been found in general population samples.[1–9,15,17] The final model suggests that individuals who have low self-esteem (measured a year prior), or who experience negative life events, or who use methaqualone, are at elevated risk for developing depressed mood. An issue to consider was the effectiveness of the nested case-control design in controlling for Time 1 depressed mood. We examined this by re-analyzing all models reported in Tables 1–4 and the final model. While the Time 1 measure of depressed mood was a strong predictor of the same variable at Time 2, the same findings emerged from the reanalysis with only slight and trivial changes in the relative risk estimates. Thus, sufficient control for prior depressed mood appears to have been achieved in the design alone.

In this sample, methaqualone use appears to substantially increase the risk for depressed symptoms independently of self-esteem or negative life events. The majority of methaqualone users who became depressed at Time 2 may have initiated their use of this substance during the study period: Of the 10 individuals with depressed symptoms who reported use of methaqualone at Time 2, seven (70 percent) had not reported use at Time 1 while three (30 percent) reported use at both time points. This methaqualone-depression association cannot be explained by the potential confounding effects of other drug use assessed in this study or other predictive variables. In fact, the adjusted relative risk estimates for methaqualone use were higher than the crude estimate. Nevertheless, the existence of a third variable which causes both a propensity to use methaqualone and to become depressed cannot be ruled out. When examined with other drug variables, heroin use was also related to depression. However, when considered together with psychosocial variables, heroin use failed to retain its importance in predicting an episode of depression.

Our review of the literature indicates that the association found between methaqualone use and depressed mood in this drug using sample is a new finding and, therefore, deserves further comment. Methaqualone is a synthetic central nervous system depressant compound that is similar to barbiturates in its sedative and hypnotic action.[38,39] Until January 1984, when it was made a controlled substance,[40] the drug was marketed in the United States under the trade names Quaalude, Sopor, Parest, Optimil, and Somnafac.[41] Its use reached epidemic proportions in the 1970s and early 1980s [42–46]; perhaps, due in part to its purported aphrodisiac qualities.[47] Although its use today has diminished as a result of discontinuation of legal production in the United States,[48] quantities produced by clandestine domestic or small foreign laboratories are still available.[47,48]

Methaqualone users are reported to be more likely to also use other illicit and licit psychoactive substances than are non-users of the drug.[44,49] This was confirmed in the present sample. The multi-drug involvement of methaqualone users could account for this study's finding of univariable associations between depressed mood and various other substances (marijuana, tobacco, hallucinogens) but lack of association when controlling for methaqualone use. Our results suggest the possible prudence of adjusting for the potential influence of methaqualone use when examining the relation between other drugs and depression.

A relation between methaqualone use and depressed symptomatology has been hinted at in two prior studies. Schwirian and colleagues,[49] in a study of personality characteristics of methaqualone users, noted a cross-sectional association between duration of methaqualone use and measures of "lethargy and introspectiveness." Craig and Van Natta[50] looked at cross-sectional, unadjusted associations between recent use of a variety of prescription medicines and the Center for Epidemiologic Studies' depression (CES-D) scale. Users of anxiolytics and sedatives (which included but were not limited to methaqualone) had significantly greater proportions of high CES-D scores than did non-users of either class of drugs with stronger associations for females than for males. Barbiturates, a class of drugs with a similar sedative-hypnotic action to methaqualone, have also been linked to depression. Recent case reports indicate that phenobarbital may cause severe depression in patients taking the drug to control seizure disorders.[51,52] At a more general level, Judd and Grant[53] found that central nervous system depressant users, especially sedative users, displayed pronounced neuropsychological deficits on cognitive tasks which assessed abstract thinking, nonverbal learning, eye-hand coordination, and perceptual accuracy.

The temporal ordering of methaqualone use and depressed mood is important in terms of etiology. In a survey of college student methaqualone users that is relevant to this issue, Kochansky and colleagues[44] report that while no particular emotion encouraged the use of methaqualone, "moods named as clearly aversive to use were depression and anger." This study suggests that use of methaqualone is more likely to be an antecedent of depressed mood than a consequence. Until more conclusive research is conducted, it would be premature to infer that methaqualone use causally contributes to the development of depressed mood. Also, it should be noted that methaqualone is an infrequently used drug. Therefore, despite the almost four-fold relative risk for depressed mood, its attributable risk both in the general population and among drug user samples would be less than that of other explanatory variables including other potential etiologic factors examined in this study.

In this sample, lower self-esteem was associated with the development of a depressive episode one year later regardless of drug use behavior or the occurrence of negative life events. Likewise, the occurrence of negative life events also independently predicted depressed mood. Such findings may suggest alternative pathways to depressed mood in drug users, with those individuals low in self-esteem, or who use methaqualone, or who experience one or more stressful life events at higher risk. A combination of these risk factors would appear to increase risk in a multiplicative manner. Each one of these three variable's ability to predict subsequent depressed affect does not depend on a particular level of the other two; that is, effect modification was not detected.

A study limitation was reliance on a measure of depressed symptomatology as opposed to a diagnostic classification for depressive disorder. Nevertheless, a strong consistency exists between psychosocial risk factors for depressed symptoms and unipolar depressive disorder indicating a possible continuity between these two conditions.[54]

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European Charter on Environment and Health

Following 12 months of intense diplomatic activity and detailed technical consultations among the 32 countries of the European Region of the World Health Organization, a ministerial conference was held in Frankfurt on December 7–8, 1989, which formally approved a Charter on Environment and Health. The European Region of WHO represents 825 million people and covers an area from Greenland in the west to the Pacific coast of the Soviet Union in the east. The conference was attended by Ministers of both Health and Environment; observers were present from Canada and the United States.

Efforts are underway to secure adoption of the Charter by all sections of society in the countries of the WHO European Region.

The decision to develop the European Charter was based on a growing perception that links between the environment and health were not clearly perceived at international, national, and local levels and that insufficient attention had been given to the integration of health dimensions into policy decisions in such fields as physical planning, energy, transport, and agriculture, nor were potential health consequences always adequately addressed at the planning stage of new developments.

The conference not only addressed specific biological, chemical and physical hazards but placed considerable emphasis on the importance of a salubrious overall environment in terms of housing, the community, and the workplace. In brief, the Charter:

- Acknowledges the dependence of human health on a wide range of environmental factors;
- Devotes a section to entitlements and responsibilities, i.e., of individuals as well as public and private bodies;
- Encompasses principles for public policy—including new policies, technologies and developments—that safeguard the environment and human health;
- Enumerates strategic elements essential to its objectives, i.e., control measures, standards, information systems, and the appropriate use of fiscal/administrative/economic instruments and land-use planning;
- Delineates priority areas to which governments, the European community, and intergovernmental organizations should pay particular attention;
- Lays down steps necessary to reverse negative trends and maintain and increase improvements already taking place, in a section titled "The Way Forward."

A copy of the Charter, as well as a detailed commentary and other relevant materials, may be obtained from the WHO Regional Office of Europe, 8 Sherfigsvej, 2100 Copenhagen, Denmark.

HIV Antibody Testing among Adults in the United States: Data from 1988 NHIS

ANN M. HARDY, DRPH, AND DEBORAH A. DAWSON, PHD

Abstract: Data collected from 21,168 adults using the 1988 AIDS supplement to the National Health Interview Survey were examined to determine awareness of and experience with HIV antibody testing in the United States. Three-fourths of adults knew of the blood test for HIV antibodies; awareness was lower among Blacks, Hispanics, older adults, and those less educated. Overall, 17 percent of adults had been tested; of these, 73 percent because of blood donation, 14 percent through other non-voluntary programs (such as military induction), and 16 percent sought testing voluntarily. While a smaller proportion of Black and Hispanic adults had been tested, they were more likely than their White non-Hispanic counterparts to have been tested voluntarily. Persons who reported belonging to groups with high-risk behaviors were also more likely to have been voluntarily tested. Most of those tested voluntarily received their test results, but only one-third also received prevention information. Three percent of adults plan to be tested voluntarily in the next year; about half will seek testing through their doctor or health maintenance organization. (*Am J Public Health* 1990; 80:586–589.)

Introduction

Counseling and testing for antibodies to human immunodeficiency virus (HIV) are important parts of the US Public Health Service plan for the prevention and control of HIV infection and AIDS (acquired immunodeficiency syndrome).[1] Throughout the United States, confidential and/or anonymous counseling and testing sites have been established to provide these services. Between 1985 and September 1988, over 1.4 million persons were counseled and tested in publicly funded test sites.[2] HIV antibody testing may also be obtained at a variety of other health care settings, but the number tested in these settings is not known. In some instances, such as blood donation and physical examination for military induction, individuals are routinely screened for HIV antibodies.

It is important to have information on the number of persons tested under all the above mentioned circumstances to help evaluate the extent to which these particular prevention strategies have been implemented. To address this need, data were examined on the adult population's experience with the HIV antibody test from the National Health Interview Survey of AIDS Knowledge and Attitudes.

Methods

The National Health Interview Survey (NHIS) is a cross-sectional survey of the US civilian, non-institutionalized population, conducted annually since 1957.[3] Each year, members of 40–50,000 households across the country are interviewed in person by Census Bureau personnel. In addition to basic health and demographic information which is collected every year, supplemental surveys on special health topics which change periodically are also included.

Since August 1987, the NHIS has included a set of questions on AIDS. Data collection with the version of the AIDS survey analyzed in this report began in May 1988 and continued throughout 1988 and 1989. The questionnaire, administered to one randomly selected adult 18 years or older in each family included in the NHIS sample, was designed primarily to estimate adults' knowledge about AIDS, HIV transmission, and prevention of HIV infection. The information obtained has been summarized in previous reports.[4–10] Questions on knowledge about and use of the HIV antibody test are also included in the survey[4–9] and form the basis of this report.

The results are based on interviews conducted from May through October 1988 with 21,168 adults or about 90 percent of eligible respondents. The data were weighted to provide national estimates. These estimates are provisional because the 1988 AIDS records were processed before the other NHIS records and have not yet been merged with them. However, the estimates should differ only slightly from those which will be made using all of the 1988 data. Provisional standard errors for these estimates were calculated by applying a design effect of 1.3 to the standard errors obtained by assuming a simple random sample. These provisional estimates of variance were used to determine the significance of differences observed between proportions; all the differences noted are significant at the 0.05 level.

For analytic purposes, persons tested for HIV antibodies included those who reported they had had their blood tested and those who did not report being tested but who had donated blood since March 1985, when routine screening for HIV antibodies in donated blood began. Information on HIV antibody testing was examined in various demographic subgroups. In addition, two other groups were examined based on information obtained in the survey. The first were persons who reported belonging to one or more of the groups with behaviors associated with an increased risk for acquiring HIV infection (such as men who have had sex with other men or persons who have used illegal drugs by needle). This question is phrased similarly to that used by the Red Cross to screen potential blood donors. Respondents were given a list of statements about risk behaviors and were asked only to indicate if at least one was true for them; they did not indicate which were true. Two percent of respondents reported that at least one of these statements was true. The second group examined included persons who felt they had a high or medium chance of having or getting HIV infection; 3.5 percent of adults fell into this category.

Results

Overall, 75 percent of adults in the US had heard of the HIV antibody test (Figure 1). The percent who were aware of the test varied by age, race/ethnicity, and education, with lower levels of awareness noted among those over 50 years

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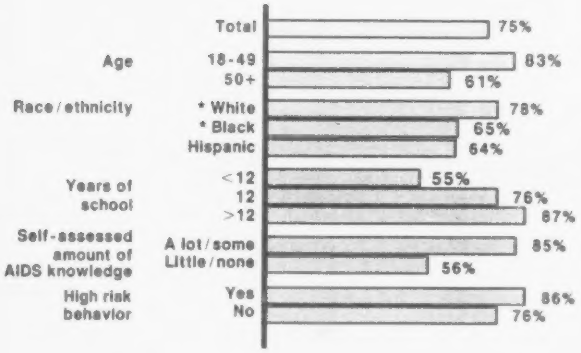


* Non-Hispanic ethnicity
SOURCE: NCHS, National Health Interview Survey, 1988

FIGURE 1—Percent of Adults Who are Aware of the HIV Antibody Test, by Demographic Characteristics

of age, among Black and Hispanic adults, and among those with less than 12 years of education. People who were more knowledgeable about AIDS and those who reported belonging to one or more of the high-risk behavior groups were more likely to know about the test.

Seventeen percent of adults had been tested for antibodies to HIV (Table 1); the majority because of blood donation since March of 1985. Men were more likely to have been tested than women, Black and Hispanic adults were slightly less likely than White non-Hispanic adults to have been tested. However, Black and Hispanic persons who were tested were more likely to have voluntary testing and testing by other non-voluntary means than their White non-Hispanic counterparts. The group with the highest proportion tested (35 percent) were those who reported high-risk behaviors for AIDS/HIV and voluntary testing was the most common circumstance for this group.

TABLE 1—Percent of Adults Undergoing HIV Antibody Testing and Distribution of Those Tested by Circumstance

Group	Percent Tested	Circumstance of Testing for Those Tested*		
		Blood Donation	Voluntary	Other non-voluntary
			Percent	
Total	17	73	16	14
Sex				
Male	20	72	15	16
Female	13	74	17	10
Race/Ethnicity				
White, non-Hispanic	17	76	14	12
Black, non-Hispanic	14	55	26	23
Hispanic	14	58	22	23
Age (years)				
18–29	24	70	17	17
30–49	20	74	16	12
Over 49	7	75	12	10
Region				
West	17	66	21	15
All others**	16	74	14	14
High-risk Behavior				
Yes	35	40	46	14
No	16	74	14	14
Chance of HIV Infection***				
High/medium	24	51	27	23
Low/none	17	74	15	13

*Multiple responses may sum to more than 100%.
**No differences between other regions (Northeast, South, Midwest).
***Chance of having or getting HIV infection.

The characteristics of those tested voluntarily, those tested by all other means (including blood donation), and those not tested are summarized in Table 2. Persons who were tested, regardless of the circumstance were more likely to be younger and better educated than those not tested. Persons tested voluntarily were more likely to be Black and to report high-risk behaviors and a greater chance of HIV infection than those tested under other circumstances and those not tested.

Most (73 percent) of those tested voluntarily were tested only once and 82 percent reported receiving results of their testing. However, of those who got their results, only 32 percent also reported receiving information on how to avoid getting or transmitting HIV.

Six percent of adults planned to be tested in the next year; 45 percent of these had not been tested previously. About half of those with such plans indicated that they would seek testing voluntarily. Figure 2 shows where those with plans for voluntary testing will reportedly go to be tested. Over half (52 percent) will get tested through a doctor or at a health maintenance organization (HMO).

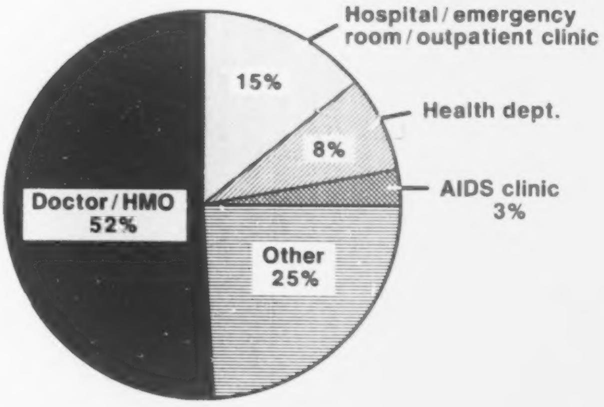
Discussion

Some caution is necessary when interpreting data from the NHIS. While the complex sampling methods used in the survey assure a representative national sample and the response rate to the survey is good, much of the data presented are based on self-report by the participants and cannot be verified. This is especially important when dealing with a sensitive topic such as HIV testing which may affect the person's willingness to respond honestly. However, the figures on HIV antibody testing obtained with the survey have remained relatively stable throughout 1988.[9] Also, the

TABLE 2—Comparison of Demographic Characteristics of Those Tested Voluntarily for HIV Antibodies, Those Tested under Other Circumstances, and Those Not Tested

Characteristics	Circumstance of Testing		
	Voluntary	Other*	Not Tested
	Percent with characteristic		
Race/Ethnicity			
White, non-Hispanic	78	85	79
Black, non-Hispanic	15	8	11
Hispanic	7	5	6
Other	5	2	3
Age (years)			
18–29	40	37	24
30–49	48	47	38
Over 49	12	16	39
Years of School			
Less than 12	14	10	24
12 or more	85	90	74
High-risk Behavior			
Yes	14	3	2
No	86	96	96
Chance of HIV Infection**			
High/medium	9	4	3
Low/none	89	95	94

*Includes those tested because of blood donation and as part of other activities which required testing.
**Chance of having or getting HIV infection.



SOURCE: NCHS, National Health Interview Survey, 1988

FIGURE 2—Distribution of Reported HIV Antibody Testing Sites for Those Planning Voluntary Testing in the Future

information on blood donation is consistent with other data on blood donors.[11-13]

In general, awareness of the HIV antibody test is high. However, it is lower than desired in certain important groups such as racial/ethnic minorities. This accounts for at least some of the deficit in testing rates among these groups; when testing rates were examined among those aware of the test, the racial/ethnic differences were smaller and no longer statistically significant. These lower levels of knowledge about the HIV antibody test and lower testing rates for Black and Hispanic adults are particularly important because of the higher rates of AIDS and HIV infection in these two populations.[14] There are also differences in AIDS rates among various subgroups within the Hispanic population.[15] We were unable to examine differences among these various subgroups because of the small numbers in our data set. The same survey instrument will ultimately be used for 20 months (May 1988 through December 1989) making it possible to examine HIV antibody testing experience within Hispanic subgroups in the future. Awareness of the test increased slightly from 1987 to 1988 among adults in the US[6]; however, it remained the same throughout 1988.[9] Efforts to increase the use of HIV antibody testing should also include ways to increase awareness about the test.

As of October 1988, approximately one in six civilian adults had been tested for HIV antibodies, primarily through blood donation. While those with high-risk behaviors who were tested were more likely to be tested voluntarily, 40 percent were tested as a result of blood transfusion. The observation that some individuals with high-risk behaviors for HIV infection have donated blood despite attempts to encourage self-deferral is not new. Further interpretation of these data is limited because the survey was not designed to address this particular issue. First, we have no information on whether any of these high-risk blood donors indicated at the time of their donation that their blood should not be used for transfusion. Thus, we cannot estimate how much of this donated blood was actually transfused. Secondly, we have no information on when these individuals engaged in high-risk behavior or on the specific dates of blood donation. For some of these donors, the high-risk behaviors may have occurred after their donation.

The characteristics of those seeking voluntary testing indicate that a higher proportion were in groups targeted for testing (e.g., those reporting high-risk behaviors) than those tested by other means or those not tested. However, many in these targeted groups have yet to be tested and most of those voluntarily tested perceived themselves to be at low or no risk of having or getting HIV infection and did not report belonging to groups with high-risk behaviors. This may be one reason why a large percentage reported not getting prevention counseling along with their test results. However, counseling is at least as important as testing in helping to prevent the spread of HIV and even low-risk individuals benefit from having such information. Also, some individuals who present for testing may be unwilling to admit risky behavior to health care providers. Post-test counseling on HIV prevention should be provided to all who are tested.

Of those tested voluntarily, 18 percent did not get the results of their tests. This figure includes those who were tested blindly and could not get their results and those who had not yet received their results when interviewed, as well as those who chose not to get their results. Presumably, most were in the latter category, this figure is similar to the 10–20 percent non-return rate reported to the Centers for Disease Control's Counseling and Testing Statistical Reporting System for persons tested at publicly funded HIV counseling and testing sites (K. Cahill, CDC, personal communication). At present the small size of the group who did not get their test results identified in our survey makes it difficult to further characterize them reliably; it will be important to do so in the future when the number of sample cases is larger.

The information on the number who plan to be tested in the coming year can be used to help ensure that the future demand for testing can be met. According to the figures obtained in the survey, approximately five million adults may voluntarily seek HIV antibody testing in the next year. Many of these persons indicated that they would go to their doctor or HMO to be tested. Evidence from other studies suggests that at least some physicians are reluctant to discuss AIDS and associated risk behaviors with their patients.[16,17] Thus, there continues to be a need to educate physicians and other primary health care providers so that they will have the information and skills necessary to provide adequate pre- and post-test counseling for their patients who request HIV testing.

Counseling and testing for antibodies to HIV will continue to be important parts of the effort to control the spread of this infection. The recent recommendation that HIV-infected individuals be immunologically monitored to identify candidates for prophylaxis against *Pnemocystis carinii* pneumonia may lead to an increase in testing.[18] The NHIS AIDS survey will continue to provide a way to monitor the national experience with HIV antibody testing. A public use data tape containing the data collected in 1987 is currently available; the 1988 data tape was to be available at the end of 1989. During 1990, the survey will be expanded to include more questions about HIV antibody testing.

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May 7 Designated as National Nurses' Day

Each year, a day is set aside to recognize the outstanding contributions made by the nursing profession. This year, May 7 has been designated "National Nurses' Day" and a nationwide observance will highlight the important work and health care provided by the more than two million registered nurses in the United States. The American Nurses' Association (ANA) and its 53 constituent nurses' associations will take part in activities to recognize the numerous contributions nurses make to the health of citizens of the US.

"Nurses Together in Caring" is the theme adopted by ANA for this year's nursing celebration. Nurses play a pivotal role in enhancing and promoting health care and the well-being of patients and family members. The US is experiencing a nationwide nursing shortage at the present time. During the next decade, more than 600,000 additional nurses will be needed. ANA and the state nurses' associations are working together to combat this shortage and misperceptions about nursing as a career. ANA is calling for organized groups to work together to solve the nursing shortage problem, help improve nurses' salaries and working conditions, and provide a greater role for nurses in decision-making processes. For additional information, contact: ANA, 2420 Pershing Road, Kansas City, MO 64108.

NURSES

Together in Caring

National Nurses' Day • May 7, 1990

Persistence of Anti-HBs among Health Care Personnel Immunized with Hepatitis B Vaccine

MARY T. PASKO, PHARM D, AND THOMAS R. BEAM, JR., MD

Abstract: Health care personnel who received the hepatitis B vaccine (Heptavax-B^R, MSD) were followed for persistence of hepatitis B surface antibody (anti-HBs). Response occurred in 135/146 (92.5 percent) vaccinees. Loss of anti-HBs (<72 RIA units; 10 S/N) occurred in 35.9 percent during the 36-month surveillance. Stepwise discriminant analysis found age and magnitude of initial antibody level, but not weight-height index, to be predictive of antibody loss over the 36 months. Twenty-four of 27 employees (88.9 percent) who lost anti-HBs responded to a fourth vaccine dose. In contrast, three of eight initial non-responders (37.5 percent) developed antibody after a fourth vaccine dose. (*Am J Public Health* 1990; 80:590–593.)

Introduction

Hepatitis B virus (HBV) infection is a significant occupational risk to health care personnel.[1-3] Safety and immunogenicity of the hepatitis B vaccine have been demonstrated in immunocompetent health care personnel.[4] However, current evidence from studies of healthy vaccinees and male homosexuals suggests serum antibody will not be maintained indefinitely.[5-9] Whether loss of demonstrable antibody can be equated with loss of protection against disease is not known.

The objectives of the present study were to determine prevalence of baseline antibody to hepatitis B surface antigen (anti-HBs) in our population, to evaluate response rate, and to monitor persistence of anti-HBs serially over three years. Non-responders and those responders who lost antibody (transient responders) were also evaluated and compared to the persistent responder group.

Methods

Participants were health care personnel at the Veterans Administration Medical Center, Buffalo, New York. The study was approved by the Human Studies Subcommittee and each participant signed informed consent prior to enrollment.

The clinic nurse recorded height, weight, health problems and/or medications upon entry into the study. A baseline serum sample was obtained and three doses of Heptavax-B^R (MSD) (20 mcg) were administered intramuscularly according to current recommendations; second and third doses were given one and six months after the first.[10]

Serum samples were collected 7, 12, 24, and 36 months following the first vaccine dose for determination of anti-HBs. Non-responders and those employees with demonstrable loss of anti-HBs to levels less than or equal to 72 radioimmunoassay (RIA) units were given a fourth vaccine dose (20 mcg). Fourth doses were given to non-responders two to three months after the third dose. Transient responders received fourth doses at variable intervals with respect to the third dose but within two to three months of a serum sample demonstrating anti-HBs $\leq$ 72 RIA units. Serum anti-HBs was then determined one month later. All sera from non-responders were tested for the presence of hepatitis B surface antigen (HBsAg).

Participants were asked to report any known or suspected exposure to HBV. Exposures were investigated by study personnel to determine whether a true HBV exposure occurred.

Serum Studies

Serum anti-HBs was determined in duplicate by radioimmunoassay (AUSAB, Ausria II; Abbott Laboratories, North Chicago, IL). Anti-HBs levels were quantitated according to manufacturer's directions and reported as radioimmunoassay units (RIA units).[11] In our laboratory, 72 RIA units correlated with 10 S/N units (ratio of sample counts per minute to mean counts per minute in antibody negative controls).

Vaccine response was defined as an anti-HBs value of greater than 72 RIA units (10 S/N). Those with a maximal anti-HBs value of 72 or less either 7 or 12 months after the first vaccine dose were considered nonresponders. Persistent responders were those whose anti-HBs values exceeded 72 RIA units (10 S/N) throughout the period of study. Those responders whose anti-HBs values initially exceeded 72 RIA units but declined below this value during the course of the study were considered transient responders.

Weight-height index was calculated as weight (kg)/{height (m)}^{p}, where p equals 2 for males and 1.5 for females.[12]

Statistical Analysis

Data analysis was accomplished with the use of the Statistical Package for Social Sciences (SPSS-PC) (v.2.0; SPSS Inc., Chicago, IL; 1988). Students t-test for unpaired data and stepwise discriminant analysis were performed; 95% confidence intervals (CI) for mean differences were calculated.[13]

Results

Baseline Screen for Anti-HBs

Of 184 employees screened for the presence of serum anti-HBs, 14 (7.6 percent) were found to be positive (>8 RIA units; 2.1 S/N). Anti-HBs levels in six of the 14 sera (42.9 percent) were 72 RIA units or less while eight sera (57.1 percent) exceeded this value. Those with baseline anti-HBs > 8 RIA units were excluded from further analysis.

Response to Vaccination

Of the 170 employees with negative baseline serum, 17 employees either did not complete vaccination or were lost to

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follow-up and seven employees received injections in the buttocks. All of these were excluded from further analysis.

Among the 146 employees remaining, 135 (92.5 percent) responded while 11 (7.5 percent) were non-responders. The mean age of the study population was 36.0 ± 9.9 years. Non-responders (mean 40.5 years) were slightly older than responders (mean 35.7 years; 95% CI of difference = 2.8, 12.4). The mean calculated weight-height index for responders (29.6 ± 6.3) and non-responders (28.9 ± 5.3) was similar.

Persistence of anti-HBs

Antibody concentrations declined to 72 RIA units (10 S/N) or less in 6 of 133 employees (4.5 percent) at 12 months, 21 of 115 (18.3 percent) at 24 months, and in 33 of 103 (32.0 percent) at 36 months after the first vaccine dose (Figure 1). The duration of anti-HBs was related to the magnitude of the antibody response measured seven months after the first vaccine dose (Table 1).

To further characterize the persistent and transient responders, 92 individuals were evaluated from whom serial anti-HBs values were available 7, 12, 24 and 36 months after the first vaccine dose. These 92 individuals were similar with respect to age, initial anti-HBs level and weight-height index when compared to the 43 responders for whom follow-up was incomplete. Anti-HBs levels were maintained above 72 RIA units (10 S/N) in 59 individuals (64.1 percent) for the 36-month period, while 33 (35.9 percent) were transient responders.

Stepwise discriminant analysis revealed that magnitude of the initial anti-HBs concentration and age were significant discriminators of antibody persistence or loss within the 36-month period. Transient responders were older than persistent responders and the magnitude of initial anti-HBs was substantially higher among persistent responders. The weight-height index (WT-HT) was not found to be of predictive value (Table 2).

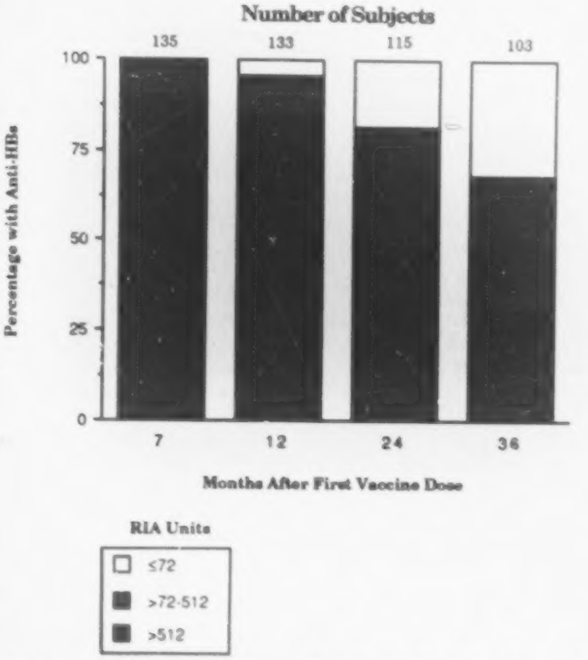

FIGURE 1—Persistence of Anti-HBs over Three Years in Health Care Personnel Who Responded to the Three-dose Vaccine Series

TABLE 1—Loss of Anti-HBs 36 Months after Vaccination (N = 103)

Initial Anti-HBs[a]		No. < 10 S/N /total (%)	Duration of Anti-HBs (months)[b]
RIA Units	S/N		
72–512	10–50	11/11 (100)	16.3 ± 6.2
>512–5000	50–1000	14/21 (66.7)	25.5 ± 7.1
>5000	>1000	8/71 (11.3)	32.6 ± 3.4

[a] measured 7 months after first vaccine dose.
[b] mean ± standard deviation of time (months) that anti-HBs ≥ 10 S/N.

Exposure to HBV

Twenty-five employees reported 110 exposures to potentially HBsAg positive blood or body fluid. Exposures involved direct contamination of mucous membranes or abraded skin with blood or other body fluids; none was an accidental needle stick. Antigen status of the patients was determined or known in 104 of the 110 (95 percent) incidents and 30 (29 percent) involved HBsAg positive patients. The 30 documented exposure events occurred among nine employees, two of whom were non-responders. None of the nine employees developed HBsAg.

Revaccination of Non-responders and Transient Responders

Response to a fourth vaccine dose was assessed in 27 transient responders and eight non-responders. In comparison to the response elicited following the initial three-dose series, anti-HBs following the fourth dose was of similar or greater magnitude in 24 of 27 (88.9 percent) transient responders. Anti-HBs exceeded initial magnitude by three-fold or greater in 13 individuals (48.2 percent), was of similar magnitude in 11 (40.7 percent), and declined in two individuals (7.4 percent). In one case (3.7 percent), the fourth dose failed to evoke any anti-HBs response. However, this individual initially demonstrated a low antibody response to the three-dose series (anti-HBs = 112 RIA units; 18 S/N). Persistence of anti-HBs following the fourth dose in this subpopulation is under continued evaluation.

In contrast, five of eight non-responders (62.5 percent) given a fourth dose failed to respond. Among the three individuals who responded, anti-HBs persisted above 72 RIA units (10 S/N) for 9, 27, and 36 months, respectively.

Discussion

The results of this study corroborate the findings of others but also provide important additional information about persistence of antibody levels and response to booster injections in recipients of plasma-derived hepatitis B vaccine (Heptavax-B). Since we found a low prevalence of hepatitis B markers among our employees (7.6 percent), these observations support the contention that antibody response and decline in our employee population was indicative of the vaccination process and not inadvertent exposures to contaminated blood and body fluids.

In the present study, initial response rate, magnitude of mean peak anti-HBs and loss of antibody within the first year were very similar to that reported by Dienstag, *et al*,[4] (Table 3), although the relation to age was not as strong as relation to initial response. Other studies[12,14,15] have found age to be an important factor distinguishing responders and nonresponders.

TABLE 2—Comparison of Persistent and Transient Responders followed Serially for 36 Months[*] (N = 92)

	Persistent	Transient	Difference (95% CI)	p-value[b]
anti-HBs[a]	29951 ± 19409	7470 ± 14450	22481 (14896, 30066)	.00001
age	33.9 ± 8.7	41.7 ± 9.7	7.8 (3.9, 11.7)	.0002
WT-HT	29.4 ± 4.8	31.1 ± 8.7	1.7 (−1.1, 4.4)	.2128

[*] data expressed as mean ± standard deviation.
[a] initial antibody level at 7 months post dose 1 expressed in RIA units.
[b] stepwise discriminant analysis.

TABLE 3—Response to Vaccination among Two Populations of Health Care Personnel

	Dienstag et al[a] (N = 645)	VAMC (N = 135)
Initial Anti-HBs[b]		
≥100,000	43.1%	NA
≥51,200	NA	23.8%
≥1,000	79.2%	79.4%
≥512	87.0%	90.4%
Mean Peak Anti-HBs[b]	26,718	26,606
Transient Responders[c]	2.3%	4.5%

[a] reference number 4.
[b] RIA units.
[c] Anti-HBs loss within 12–13 months of the first vaccine dose.

Although the influence of injection site on response rate is well established,[16] the role of obesity as a factor in response rate is less clear. In our study, no association was found between likelihood of response and weight-height index as a measure of obesity. In a previous study reporting an association,[12] all vaccine was administered in the buttocks with a 1-inch needle and thus the overall seroconversion rate was low as might be expected (55.7 percent). Horowitz, *et al*,[7] found an association between body mass index (weight/height^2) greater than 25 and prevalence of low antibody levels (<10 mIU/ml) at a median time of 39 months post-vaccination. The subpopulation with low anti-HBs (<10 mIU/ml) included non-responders as well as responders who lost antibody since vaccination. A recent case control study reported by the Centers for Disease Control (CDC) used an age-adjusted logistic model and found an association between obesity and nonresponse for females over the age of 19 years. A similar association was not found for males.[17]

Antibody persistence among vaccine responders in health care populations is of particular interest. During our 36-month surveillance, 32 percent of responders lost antibody (<10 S/N). Jilg, *et al*,[5] in a brief report of "healthy vaccinees" (inclusion groups not further characterized) found antibody in 27.5 percent of the study population fell to 10 IU/l (equivalent to 10 S/N) or below after four years. The authors do not comment on site of vaccine injection, age of their study population, or HBV exposure events.

Several additional studies have attempted to evaluate either initial response or antibody persistence following vaccination of health care personnel. However, vaccine was administered via gluteal injections in two reports[14,15] and response rate was low as expected. In an additional two reports,[9,10] initial antibody levels post vaccination (deltoid injections) were not obtained. Both authors assumed a response rate of 90–95 percent and thus the subset of employees reported to be negative for anti-HBs at the respective serum sampling times included non-responders and low level responders. Additionally, since baseline prevalence of anti-HBs is also unknown in these populations, those with persistent antibody as a result of natural infection cannot be distinguished from vaccine responders.

Extensive long-term follow-up of vaccinees has been reported by Hadler, *et al*, in a study of male homosexuals[6] and by Wainwright, *et al*, in an Eskimo population.[9] In the Hadler study, anti-HBs was maintained above 10 SRU (sample ratio units; equivalent to 10 S/N) in 93, 70, 65, 60, and 58 percent of the population at 12, 24, 36, 48, and 60 months of follow-up, respectively. Similar to Hadler's data, anti-HBs in our study was maintained above 72 RIA units (10 S/N or 10 SRU) in 95, 82, and 68 percent of vaccinees at 12, 24, and 36 months of follow-up. Preliminary review of our 48-month data indicates 52 percent of vaccinees are above 10 S/N (unpublished observation). This percentage is similar to that of Barnes, *et al*,[8] at 48 months (55 percent > 10 S/N). In contrast, Wainwright, *et al*,[9] report 73 percent above 10 S/N at five years post-vaccination for ages 20–49 and 63 percent for those over 50 years of age. Similar to our finding, initial antibody level was a stronger predictor of antibody persistence than age.

Characterization of responders who subsequently experience a loss of anti-HBs has not been previously reported. We found these transient responders to be somewhat older and have lower initial anti-HBs levels than those with persistent antibody. Evaluation of the response to booster doses and persistence of antibody is in progress. Preliminary results from a small group of our transient responders indicate that a single booster dose elicits a response equal to or greater in magnitude than the initial response in 89 percent of recipients. Continued follow-up of this group will provide further information about persistence of antibody.

While the majority of our transient responders developed anti-HBs after a fourth dose of vaccine, the need for booster doses of vaccine is controversial. Although progressive loss of anti-HBs occurs over time, immunity may outlast demonstrable antibody. Among male homosexuals in the Hadler study,[6] the risk of hepatitis B infection was greatest among vaccine non-responders or low level responders. Of the hepatitis B infections that occurred, 76 percent were in individuals whose anti-HBs was below 10 SRU (equivalent to 10 S/N). Only two cases of clinically important hepatitis developed among responders. Likewise, Wainwright, *et al*,[9] found only four cases of HBV infection over five years. All had anti-HBs level < 20 SRU (equivalent to 20 S/N), developed anti-HBc, remained negative for HbsAg, and were asymptomatic. Thus while current evidence suggests that maintenance of demonstrable antibody levels within five years of vaccination may be unnecessary for prevention of symptomatic disease, further studies are required to investigate the issue of life-long protection.

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DiMarzio, RN, for dedicated clinical support. We also thank M. Feldman, PhD, of Medical Research Service, VAMC, for assistance with statistical analysis. The study was presented in part to the 26th Interscience Conference on Antimicrobial Agents and Chemotherapy, New Orleans, Louisiana, September 1986.

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Maternal and Child Health Law Changes Explained

Recent legislative changes in the maternal and child health block grant program were spelled out in a letter sent by Dr. Robert G. Harmon, HRSA Administrator, to state health officers. The letter is intended to assist states in applying for fiscal 1991 maternal and child health block grant funds and in submitting the fiscal 1989 annual report on state activities funded under this program.

The law was amended by the Omnibus Budget Reconciliation Act of 1989 (Public Law 101-239). The amendments "create a new framework for action on maternal and child health for the 1990s," Dr. Harmon noted.

The amendments, which increase the authorization for appropriations to $686 million for fiscal 1990 and subsequent years, are designed to:

- Promote the PHS Health Objectives for the Nation for the Year 2000;
- Assure expenditure of funds for preventive and primary care services for children and for services for children with special health care needs;
- Improve data collection and analysis; and
- Increase overall accountability in the administration of the program.

The amendments also make structural changes in the block grant to improve state planning, accountability, and targeting of federal funds to priority populations. The changes:

- Prescribe new requirements for the use of allotment funds;
- Establish a new application requirement;
- Establish new state and federal reporting on services, expenditures and health status of target populations; and
- Mandate additional state and federal activities to promote improved access to maternal and child health services.

The changes reflect "work remaining to be done by federal and state authorities in developing effective statewide systems of comprehensive, community-based family-centered continuous care," Dr. Harmon wrote. "They encourage using resources better through effective coordination of MCH block grant activities with those of other major maternal and child health assistance programs providing related services to similar populations."

For further information, contact Frank Sis, HRSA, USPHS, DHHS, 5600 Fishers Lane, Rockville, MD 20857; (301) 443-3377.

Carbon Monoxide in Indoor Ice Skating Rinks: Evaluation of Absorption by Adult Hockey Players

Benoît Lévesque, MD, Eric Dewailly, MD, MSc, Robert Lavoie, MD, Denis Prud'Homme, MD, MSc, and Sylvain Allaire, BSc

Abstract: We evaluated alveolar carbon monoxide (CO) levels of 122 male, adult hockey players active in recreational leagues of the Quebec City region (Canada), before and after 10 weekly 90-minute games in 10 different rinks. We also determined exposure by quantifying the average CO level in the rink during the games. Other variables documented included age, pulmonary function, aerobic capacity, and smoking status. Environmental concentrations varied from 1.6 to 131.5 parts per million (ppm). We examined the absorption/exposure relationship using a simple linear regression model. In low CO exposure levels, physical exercise lowered the alveolar CO concentration. However, we noted that for each 10 ppm of CO in the ambient air, the players had adsorbed enough CO to raise their carboxyhemoglogin (COHb) levels by 1 percent. This relationship was true both for smokers and non-smokers. We suggest that an average environmental concentration of 20 ppm of CO for the duration of a hockey game (90 minutes) should be the reference limit not to be exceeded in indoor skating rinks. (*Am J Public Health* 1990; 80:594–598.)

Introduction

Over the past few years numerous cases of carbon monoxide (CO) intoxication in indoor skating rinks have been described in medical literature.[1–9] Most often, the victims have been hockey players, and the source of the contamination has been the ice resurfacing machine.

When inhaled, CO quickly enters the bloodstream where once having displaced the oxygen, it becomes associated with hemoglobin to form carboxyhemoglobin (COHb). The percentage of COHb is the main indicator of the absorption of this toxic gas. It is generally equal to 1 percent among non-smokers[10] and may reach 5 to 10 percent among smokers.[11]

Following acute exposure to CO, there is a rapid increase in the COHb level. At the 3 to 4 percent level, the first signs of adaptation in healthy individuals take the form of an alteration of alertness[11] and arterial dilation in response to a drop in blood oxygenation.[12] These levels are sufficient to worsen coronary symptoms among persons with cardiac problems.[13–16] At 4 percent, careful driving skills are affected,[17] and, between 5 and 10 percent, visual acuity is decreased.[12] Levels between 10 and 20 percent cause headaches and difficult breathing.[12] At 20 percent, headaches worsen, nausea appears and manual dexterity is affected.[12] At 40 to 50 percent, all symptoms intensify, and between 50 and 60 percent, there is a possibility of coma.[12]

Gasoline or propane-driven ice resurfacers give off combustion by-products in their exhaust fumes. Depending on how well the unit operates, there will be a variable level of CO emitted. The problem is generally aggravated by poor ventilation in indoor rinks. The combination of these two factors often results in high levels. The CO reaches its maximum concentration when the ice resurfacer is in operation and gradually drops afterwards.[18]

Extremely high levels of CO have been documented in skating rinks with some levels exceeding 100 parts per million (ppm).[1,3,6,8,9,18–21]

CO absorption is up to three to four times higher during exercise than during rest.[22] We conducted a study among male adult hockey players of the Quebec City area to evaluate the CO absorbed by the players at the end of a 90-minute game. The data were collected between October 26, 1988 and January 28, 1989.

Methods

We chose 10 of the 16 voluntary hockey teams, each playing in its own indoor rinks, guiding our choice on various indicators such as the age and size of the rink, the type of resurfacing machine, and whether or not there was a mechanical ventilation.

Measuring Absorption

To measure absorption, subjects, standing up, took a deep breath which they held for 20 seconds. They then exhaled first into one polyvinyl chloride 900 mL bag and then into two other bags joined by plastic tubing (Tygon). The second sample filled with alveolar air was analyzed within 12 hours using a Bendix 8501-BA infra-red CO analyzer. The quantification technique was developed beforehand, and the correlation between the CO concentration in the alveolar air in ppm and the percentage of COHb was established using a sample of 24 volunteers (13 smokers and 11 non-smokers).[23] The square of the linear correlation coefficient (r^2) was .99, the intercept was −.003 ppm with a standard error of .005 ppm, the slope was 5.39 ppm (5.4 ppm of alv CO ≈ 1% COHb) and the standard error of the slope was .13 ppm.

For each player, the alveolar CO was measured within 15 minutes before and after the game. The difference between the two results represented the CO absorbed by the subject during the match in ppm (ΔalvCO).

Measuring Exposure

Exposure was measured by collecting seven samples of air at 15 minutes intervals, from the players' arrival on the ice rink to their departure. The samples were taken from the penalty box located 1.2 meters above the ice using a Gilian HFS-113A pump and Calibrated Instrument 5 L aluminized bags. Pump flow was set at one liter per minute. These samples were analyzed within 12 hours on the Bendix spectrometer. The average of the seven samples was con-

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sidered the level of exposure by players during the game ([CO]).

Other Variables

Player's age and whether a smoker or non-smoker were obtained by questionnaire. The pulmonary function was quantified with a Roxon Compact vitalograph using the standardization criteria of the American Thoracic Society.[24] We determined the forced expiratory volume in one second (FEV_1), the forced vital capacity (FVC), the mean forced expiratory flow (FEF_{25-75}) and the FVC/FEV_1 ratio. The expected values were estimated using the Knudson method.[25] Finally, players' aerobic capacity ($\dot{V}O_2$max) was evaluated using a sub-maximum progressive effort test on an ergometric bicycle. The initial work load was set on the basis of the weight of the individual and increased progressively by 240 kiloponds meter (Kpm) for four two-minute levels. The $\dot{V}O_2$max was estimated based on the equation suggested by the American College of Sports Medicine (26).

Statistical Analysis

The data were stratified according to smoking habits (smokers vs non-smokers). In addition to CO exposure, we examined the influence of age, $\dot{V}O_2$max and the various spirometric indices on absorption, in a multiple regression model using a backward elimination procedure.[27] The pre-selected critical value of the F distribution was set for a significance level of .05. The resulting relationship between absorption and exposure was evaluated with a simple linear least squares regression model using individual values. When heteroscedastic, it was weighted by the inverse of the variance of the means to confirm the validity of the statistical tests.[27] To verify the possible parallelism between smokers and non-smokers, the coefficients were compared with the addition of a term of interaction.

Results

We evaluated 122 players in 10 teams, 32 smokers and 90 non-smokers. Their age varied between 19 and 59 years old (mean = 34.04 years, standard deviation = 7.10 years).

The various spirometric parameters (FEV_1, FVC, FEF_{25-75}, FVC/FEV_1) as well as the age and the $\dot{V}O_2$max were eliminated by the backward elimination procedure (data available on request to authors). It seems that these variables have very little effect on CO absorption, whatever the smoking habits. As a result, these variables were not considered factors likely to alter the exposure/absorption relationship.

Tables 1 and 2 describe by indoor rink, for non-smokers and smokers respectively, exposure concentrations in ppm, the number of players, the average quantity of alveolar CO (mean and variance) absorbed during the game and the estimated equivalent in percentage of COHb (5.4 ppm alv CO ≈ 1% COHb).

The exposure levels varied from 1.6 to 131.5 ppm. Seven rinks were below 20 ppm, whereas three exceeded 75 ppm.

Table 1 shows that with low exposures (rinks 1 and 3), non-smokers had, at the end of their game, lowered their COHb concentration. In seven other rinks there was absorption and it was proportional to the level of exposure. The concentration of 4 ppm noted in rink 4 represented a balance threshold where there was neither absorption nor elimination.

TABLE 1—Variation in the Alveolar Absorption of Carbon Monoxide among Players during a 90-minute Hockey Game (ΔalvCO) and the Estimated Equivalent in Percentage of Carboxyhemoglobin (COHb) based on the Concentration of Carbon Monoxide ([CO]) in the Indoor Skating Rinks among Non-smokers

			ΔalvCO Mean Variance		
Arena	Players N	[CO] (ppm)	(ppm)	(ppm^2)	COHb %
1	8	1.6	−.46	.10	−.09
2	12	9.3	3.48	.46	.66
3	13	1.6	−.24	.10	−.04
4	9	4.0	−.04	.44	−.007
5	10	9.5	2.39	.93	.45
6	5	15.1	4.51	1.56	.85
7	14	131.5	69.79	128.90	13.17
8	8	107.3	57.36	35.07	10.82
9	5	16.2	7.81	.70	1.47
10	6	79.3	46.10	40.42	8.70

TABLE 2—Variation in the Alveolar Absorption of Carbon Monoxide among Players during a 90-minute Hockey Game (ΔalvCO) and the Estimated Equivalent in Percentage of Carboxyhemoglobin (COHb) Based on the Concentration of Carbon Monoxide ([CO]) in the Indoor Skating Rinks among Smokers

			ΔalvCO Mean Variance		
Arena	Players N	[CO] (ppm)	(ppm)	(ppm^2)	COHb %
1	3	1.6	−6.38	46.32	−1.20
2	2	9.3	−27.72	20.58	−5.23
3	3	1.6	−10.68	11.30	−2.00
4	1	4.0	−18.13	—	−3.42
5	1	9.5	−13.79	—	−2.60
6	4	15.1	−2.38	31.42	−0.45
7	6	131.5	51.50	299.45	9.72
8	6	107.3	45.66	98.87	8.61
9	4	16.2	−9.93	1.81	−1.87
10	4	79.3	34.80	357.27	6.57

Table 2 shows an average drop in COHb among smokers at exposure levels as high as 15.1 and 16.2 ppm. This phenomenon is attributable to high basic levels among smokers.

Figures 1 and 2 show the relationship between exposure and absorption using a simple linear regression model.

Figures 1 illustrates this model among non-smokers. The r^2 was equal to 0.97 and the intercept to −1.72 ppm (standard error = .69 ppm). This small negative intercept reflected the CO elimination in low exposure concentrations for the non-smokers. The slope at 0.55 ppm (standard error = .01 ppm) indicated that for every 10 ppm of exposure, an adult hockey player (non-smoker) had, at the end of a 90-minute game, absorbed enough CO to raise his alveolar CO concentration by 5.5 ppm (≈1% of COHb). The F ratio was 2681.36 for a significance level below 0.0001. The test had to be interpreted with caution given the heteroscedasticity of the data. Indeed, there was an increase in the variability of absorption proportional to the level of exposure (see Figure 1 and Table 1). For exposures of 1.6 and 131.5 ppm, absorption varied respectively from −0.9 ppm to 0.1 ppm (range = 1 ppm, variance = 0.10 ppm^2) and from 51.1 ppm to 87.4 ppm (range = 36.3 ppm, variance = 128.90 ppm^2). To correct this situation, the results were weighted with the inverse of the variance of the means. The results of the regression and the test continued to be very significant ($F = 1454.70$, $p < 0.0001$).

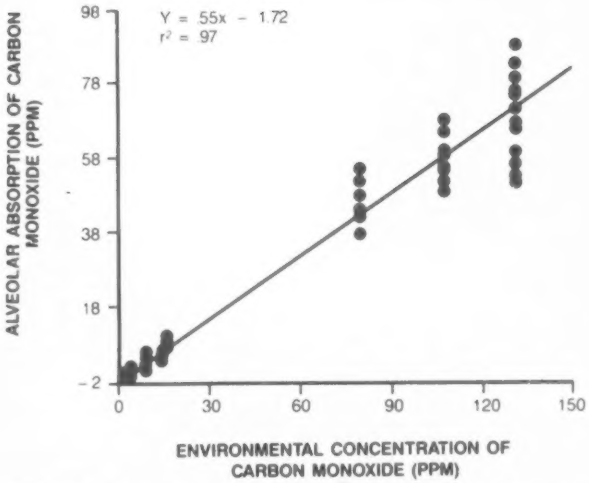

FIGURE 1—Relationship between Alveolar Absorption of Carbon Monoxide among Players during a 90-Minute Hockey Game based on Carbon Monoxide Concentrations in Indoor Skating Rinks (non-smokers).

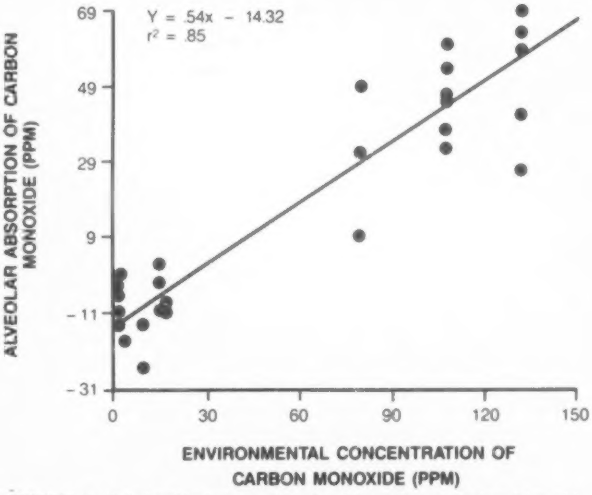

FIGURE 2—Relationship between Alveolar Absorption of Carbon Monoxide among Players during a 90-Minute Hockey Game Based on Carbon Monoxide Concentrations in Indoor Skating Rinks (smokers).

Among smokers (Figure 2), the r^2 was equal to 0.85 and the intercept to −14.32 ppm (standard error = 3.1 ppm) indicated the effect of exercise on CO elimination with elevated basic levels of alveolar CO. The F ratio was 169.50 ($p < 0.0001$). The slope at 0.54 ppm (standard error = .04 ppm) was not considered statistically different to that of non-smokers ($p = .6607$). As a result, after a 90-minute game of hockey, smokers and non-smokers alike increased their concentration of COHb by 1 percent for every 10 ppm of exposure to CO in the indoor rink.

Discussion

There are two types of standards for CO: standards for workers and standards for the public in general.

In an occupational environment the National Institute for Occupational Safety and Health (NIOSH) recommends a maximum time-weighted average (TWA) of 35 ppm for an eight-hour work day and a ceiling concentration of 200 ppm.[10] In the province of Quebec (Canada), the prevailing regulation stipulates 50 and 400 ppm for the same parameters.[28]

These values cannot apply to a hockey player playing in a recreational league. The latter, besides practicing his sport during his leisure time, is involved in an activity that required an energy expenditure greater than the workers. For example, the average energy expenditure of professional activities like lumberjack, mason, miner or carpenter is between 5.0 and 6.4 Mets (1 Met = 3.5 mL $O_2 \times Kg^{-1} \times min.^{-1}$).[29] For a hockey player, it is 12.9 Mets (min. = 9, max. = 15), a greater quantity than a soccer player (10.3 Mets, min. = 8, max. = 12) or a basketball player (11.1 Mets, min. = 9, max. = 13).[29] It is therefore necessary to look to criteria that apply to the general public.

Table 3 gives a summary of Canadian[30] and American[31] standards and of the recommendations of the World Health Organization (WHO).[32] These levels have been deduced using empirical equations based on the controlled exposure of volunteers. Apart from the desirable Canadian standard, they aim at maintaining the COHb concentration for non-smokers at 2 percent, thus protecting the most vulnerable segment of the population from the harmful effects of CO. The desirable Canadian standard aims at allowing an average COHb content below 1 percent, namely the equivalent of the endogenic production of a non-smoker.[30]

In indoor rinks, the British Columbia Department of Labor (Canada) has proposed a reference limit of 25 ppm[8] and, following a review of the literature on the subject, Luckurst and French suggested CO concentrations of 25 and 12 ppm for one hour and eight hours, respectively.[33]

The study that we conducted allowed us to ascertain extremely high CO concentrations in three indoor rinks out of 10. These levels even exceeded the standards set for workers.

Beyond environmental measurements, the best way to describe the risk involves using biological measurements. Only one author has shown interest in this aspect. Spengler evaluated alveolar breath samples among 12 Harvard University players before and after 90 minutes of exercise. An average concentration of 22.5 ppm of CO in the air was sufficient to raise the average COHb level by 1.1 to 3.2 percent.[18]

We succeeded in determining the absorption curve of adult hockey players who practice their sport on a recreational basis, and we noted that absorption was directly related to exposure.

There was great individual variability of absorption at the same CO exposure. For the non-smokers, at low level of CO, the intrateam variability was very low (variance = .316 ppm^2 for teams 1 and 3). At higher exposures, the observed

TABLE 3—Recommended Carbon Monoxide Levels for Ambient Air

	Average per 1 hour (ppm)	Average per 8 hours (ppm)
WHO* (1979)	25	13
United States (1985)	35	9
Canada (1987)		
Desirable	13	5
Acceptable	30	11

*World Health Organization

variability may be explained in part by the analytical error on the measure of alveolar CO itself.[23] Also, other variables undocumented here can probably affect the individual CO absorption at these levels. The ratio of the tidal air to the dead space of the lung and the diffusion constant of the lung can explain some divergencies in the CO uptake.[22] Moreover, factors which affect the players' activity may be important. For example, Green showed differences in the time-motion characteristics (actual playing time, number of playing shifts, recovery time between shifts, etc.) between players depending on the position played (center, wing, defense).[34] In addition for the smokers, the variability of the pre-exposure alveolar CO level is certainly an important factor.

In the more exposed teams (#10-8-7), the average increase of COHb were respectively 8.70 percent, 10.82 percent, and 13.17 percent. Some players had an increase up to 17 percent. Levels like this are certainly sufficient to cause headache, breathing difficulties and, for some people, nausea.[12] They can also alter the choice reaction time,[11] the driving skills,[17] and the visual acuity.[12] All these toxic actions can contribute to accidents on the skating rinks and on the road at the end of the game.

At the present time there are no reliable data on cardiac pathology among hockey players. However, sudden increases in COHb as noted here may have a significant impact on the triggering of an irreversible ischemia. Particularly, if they are within the framework of intense physical exercise. Montgomery estimated for 12 adult recreational players that the intensity of hockey play assessed by the Karvonen method was 85 percent.[35] This represents a strenuous activity and, by itself, it can strain the cardiovascular system.[35] All these considerations make CO a significant threat for the population under study.

Generally, these players drop out of hockey several years and then return to improve their physical fitness or for the friendship obtained.[35] They are often ill-prepared for such intense physical exercise and, of all the groups frequenting indoor rinks, they are probably the most susceptible to the harmful effects of CO given the high level of absorption conferred by the intensity of their activity. If a standard is to be issued, we believe that they should serve as a reference.

We find absurd the idea that a hockey player must adapt to a toxic gas. A 3 percent COHb threshold, which is sufficient to induce physiological compensation mechanisms, must be the reference index. Considering that non-smokers have COHb levels of 1 percent and less, an increase of more than 2 percent cannot be accepted during the hockey game. This corresponds, for the duration of the hockey game (90 minutes), to an average concentration of 20 ppm in the air of the indoor rink, and as a result, it is the limit we propose.

If we rely on the data of our study, this norm is realistic since seven indoor rinks in 10 fell within it. When an ice resurfacing machine with a non-polluting motor (electrical, hydrogen) is not available, there are three means of respecting this norm: good maintenance of the ice resurfacing machine, sufficient mechanical ventilation used wisely, and routine monitoring of CO levels in indoor skating rinks.[7,8,20,36]

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AADS to Offer Oral Health Education for Elderly

The American Association of Dental Schools (AADS) is preparing an education packet to help older persons take better care of their oral health. Funding for the program is being provided by a $49,479 grant from the W. K. Kellogg Foundation.

Through the AADS project, copies of the oral health education packets will be distributed to 55 dental schools and 200 dental hygiene education programs nationwide. Beginning in the Fall 1990, the dental and dental hygiene students will give one-hour presentations at senior centers, clubs, churches, and volunteer organizations, with a special emphasis placed on reaching minority older persons.

The educational materials will be prepared by a team of dentists, dental hygienists, a social worker, and dental educators. By conducting the training, the students will gain first hand experience about oral health and its relation to the aging process, as well as learn to communicate effectively with elderly persons and to network with other health professionals.

For more information, contact the American Association of Dental Schools, 1625 Massachusetts Avenue, Suite 502, Washington, DC 20036; (202) 667-9433.

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Public
Health Briefs

Passive Smoking and 20-Year Cardiovascular Disease Mortality among Nonsmoking Wives, Evans County, Georgia

CHARLES HUMBLE, MS, JANET CROFT, MPH, ANN GERBER, MSPH, MICHELE CASPER, MSPH, CURTIS G. HAMES, MD, AND HERMAN A. TYROLER, MD

Abstract: The association of passive smoking and cardiovascular disease (CVD) mortality was assessed in a cohort of 513 rural, married Black and White women who were disease-free and self-described as never-smokers at baseline in 1960. Over a 20-year period, 76 of 147 total deaths were attributed to CVD. Relative risk estimates adjusted for age, cholesterol, blood pressure, and body mass from proportional hazards models were 1.59 for CVD (95% CI = 0.99, 2.57) and 1.39 (CI = 0.99, 1.94) for all cause mortality among women with husbands who smoked cigarettes. (*Am J Public Health* 1990; 80:599–601.)

Introduction

Cardiovascular diseases account for about one-half of all deaths in the United States annually.[1] Although active smoking is well-established as a CVD risk factor,[2] the risk for all CVD mortality associated with passive smoking among nonsmokers has not been previously investigated. Recent studies of risks for coronary heart disease,[3–8] stroke,[4,5] or all cause mortality[7,9,10] associated with passive smoking generally have reported weak and/or statistically nonsignificant results.

The 20-year mortality experience of nonsmoking women in Evans County, Georgia was used to assess the association of passive smoking with CVD and all cause mortality. This is the first report that includes data on both Blacks and Whites and on the consistency of self-reported smoking behaviors over time.

Methods

In 1960–61, 92 percent of all residents ages 40–74 years and a 50 percent sample of individuals ages 15–39 years in Evans County, Georgia participated in a cardiovascular disease study that included risk factor measurements, complete physical examinations, and a demographic and medical history interview.[11] Detailed descriptions of the Evans County study design and the 20-year mortality follow-up of the cohort have been reported elsewhere.[11–13] At baseline, 554 (82 percent) White and 389 (83 percent) Black women, among a total of the 1,127 women ages 40–74, reported that they had never smoked. The present study was restricted to the 328 White women and 185 Black older women who had never smoked and were married to male examinees who reported they either had never smoked or were current smokers at baseline. Women married to ex-smokers were excluded from the analyses as the probability for misclassification of these subjects' own smoking habits and those of their husbands was judged to be higher than for spouses of never smokers.[14] A second survey of study subjects in 1967 provides data on the stability of reported smoking status.

Vital status was determined as of May 1, 1980. Underlying cause of death was abstracted from death certificates with codes 390-456 (ICD 8th Revision) defining CVD. All CVD mortality was chosen as an endpoint given the limitations of death certificate data and the small number of deaths attributed to each specific CVD entity.[13] Three subjects who did not have follow-up information were excluded.

Analyses for White women were stratified by social status because of its inverse relationship with smoking status and CVD mortality in this cohort.[13] White women were divided into high social status and low social status groups based on the median of the McGuire-White index of social status for all Evans County Whites. This index, based on occupation, level of education, and source of income of the head of household, was developed for use in rural settings.[15] Since only 5 percent of the Black women in the Evans County population had a social status score above the median for Whites, Blacks were not stratified by social status. Exposure to passive smoking was defined by husband's smoking status (current, never) at the time of the baseline interview.

Mean baseline characteristics by passive smoke exposure were compared using t-tests. Cox proportional hazards models[16] were used to estimate the association of passive smoking with time to all CVD, smoking-related CVD, and all cause mortality in this population while adjusting for age alone and for age, systolic blood pressure, serum cholesterol, body mass index (BMI), and a quadratic term for BMI. Relative risks (RR) and 95% confidence intervals (CI) were calculated using the SAS proportional hazards (PHGLM) modeling procedures,[17,18] and the statistical significance of trends was tested using a method proposed by Rothman.[19] Constancy of the relative risks over time was verified before the proportional hazards were modeled.

Results

Among nonsmoking married women, there were 179 (55 percent) of 328 White women and 117 (63 percent) of 185 Black women whose husbands reported current cigarette

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TABLE 1—Mean and Standard Error of Baseline Characteristics by Passive Smoking Status of Nonsmoking Wives, Ages 40–74 Years, Evans County, Georgia, 1960–61

	White Women					
	High Social Status[*]		Low Social Status[*]		Black Women	
	Exposed	Unexposed	Exposed	Unexposed	Exposed	Unexposed
N	(78)	(83)	(101)	(66)	(117)	(68)
Age	51.9 ± 1.0	54.9 ± 0.9	52.1 ± 0.8	53.9 ± 0.9	50.3 ± 0.7	55.5 ± 1.0
Systolic Pressure	145.5 ± 3.1	150.6 ± 2.9	151.6 ± 2.9	157.6 ± 4.3	170.6 ± 3.4	176.5 ± 5.0
Diastolic Pressure	88.4 ± 1.6	90.6 ± 1.4	92.2 ± 1.3	93.1 ± 1.7	103.1 ± 1.9	103.9 ± 2.5
Serum Cholesterol	231.9 ± 4.9	237.5 ± 4.5	227.0 ± 4.4	235.7 ± 7.3	216.5 ± 3.9	216.2 ± 4.6
Body Mass Index	26.3 ± 0.6	26.4 ± 0.6	27.0 ± 0.5	28.6 ± 0.9	29.2 ± 0.6	30.0 ± 0.9

[*] Based on the median of McGuire-White scores for all White subjects.

smoking behavior. Among both Black and White women there were no statistically significant ($p < 0.05$) differences by passive smoking status for systolic or diastolic blood pressure, serum cholesterol or body mass (Table 1). However, passively exposed Black women and high social status White women were younger on average than nonexposed wives by 5.2 years (95% CI = 3.0, 7.6) and 3.0 years (95% CI = 0.3, 5.5), respectively. For all Whites combined, nonexposed women were also more likely to be above the median SES (socioeconomic status) level than passively exposed women (55.7 percent vs 43.6 percent).

Comparison of self-reported smoking status in 1960 and 1967 showed 98 percent of wives again reported themselves as never having smoked in 1967. Similarly, 98 percent of never smoking husbands maintained their reported status in 1967 while 25 percent of husbands who smoked in 1960 described themselves as non-smokers in 1967.

Age-adjusted RRs for all CVD, smoking-related CVD, and all cause mortality among passively exposed wives were elevated in Blacks and high social status Whites and for all subjects combined (Table 2). The opposite relationship of mortality with passive smoking status was found for low social status White women. Adjustment for other established CVD risk factors (blood pressure, cholesterol, and BMI) generally caused modest elevations of the risk estimates (Table 3) but as with the age-adjusted estimates, the confidence intervals for all subject groups included unity. A trend in risk over level of husband's smoking as reported in 1960 was only seen among high social status Whites; RRs for both total and smoking-related CVD mortality among wives whose husbands smoked <10, 10–20, and >20 cigarettes per day as compared to wives of nonsmokers were 1.02, 2.11, and 2.55, respectively (p for trend <0.06). A marginally significant ($p < 0.09$) trend in risk for all CVD and smoking-related CVD over crude levels of duration of exposure was also apparent only among high social status White women.

TABLE 2—Age-adjusted Relative Risks and 95% Confidence Intervals for Total CVD, Smoking-related[*] CVD, and All-Cause Mortality for Wives Exposed to Passive Smoke in Evans County, Georgia, 1960–80

				Whites	
Cause of Death		All Subjects	Blacks	HSS[+]	LSS[++]
CVD Total	RR	1.34	1.69	1.66	0.60
	95% CI	0.84, 2.21	0.83, 3.46	0.64, 4.32	0.27, 1.34
Smoking-related	RR	1.29	1.57	1.67	0.61
	95% CI	0.79, 2.10	0.73, 3.37	0.64, 4.36	0.25, 1.47
All cause	RR	1.31	1.34	1.80	0.72
	95% CI	0.95, 1.82	0.79, 2.28	0.94, 3.47	0.41, 1.27

[*] ICD8 codes 410-456

[+] High social status

[+] Low social status

TABLE 3—Relative Risks[*] and 95% Confidence Intervals for Total CVD, Smoking-related[+] CVD, and All Cause Mortality for Wives Exposed to Passive Smoke in Evans County, Georgia, 1960–80

				Whites	
Cause of Death		All Subjects	Blacks	HSS[++]	LSS[−]
CVD Total	RR	1.59	1.78	1.97	0.79
	95% CI	0.99, 2.57	0.86, 3.71	0.72, 5.34	0.32, 1.96
Smoking-related	RR	1.54	1.68	1.97	0.82
	95% CI	0.93, 2.55	0.76, 3.71	0.72, 5.34	0.31, 2.15
All cause	RR	1.39	1.33	1.97	0.87
	95% CI	0.99, 1.94	0.78, 2.28	1.00, 3.90	0.48, 1.59

[*] Hazards ratios adjusted for age, diastolic blood pressure, total serum cholesterol, body mass index (BMI = kg/meter^2), and BMI^2

[+] ICD8 codes 410-456

[++] High social status

[−] Low social status

Discussion

These data suggest an elevation of risk for death from CVD and all causes among non-smoking married women whose husbands described themselves as current smokers at the beginning of a 20-year follow-up period. Our findings for Blacks are the first report associating CVD with passive smoking in this racial group. Our observations that social status may modify the effect of passive smoke exposure may be due to chance, but a similar pattern of results for coronary heart disease (CHD) has been reported in other studies of passive smoking. Nonsignificant ($p > 0.05$) two-fold RRs for CHD among passive smokers were reported from studies of middle-class and upper-middle-class women[6] and men[7] while CHD risk was significantly but more modestly increased (RR = 1.2) among a much larger sample of predominantly blue collar Washington County, Maryland women.[8] No increased risk for CHD was reported among public hospital patients whose husbands smoked in four British hospital regions.[5]

It is unlikely that these results can be explained by a change in smoking habits since the minimum age of these women in 1960 was 40. We lack data to examine whether exposure status changed during follow-up due to remarriage. The absence of elevated risk among exposed low social status

White women may reflect a failure of our passive exposure index to measure exposure within the lower social stratum. Power to test for small differences in effect of passive smoking by race or social standing was lacking as were data to evaluate the role of other variables such as alcohol use or physical activity. Taken together with the results of previous studies[6–8,10] and laboratory results suggesting that passive smoke exposure causes decreases in energy production in the mitochondria of heart muscle[20] and increased platelet aggregability in nonsmokers,[21] our results support the health hazards of exposure to passive smoke.[20]

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Community Impact of a Localized Smoking Cessation Contest

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Abstract: The present study assessed the effectiveness of a localized community contest timed to coincide with a statewide smoking cessation contest. Follow-up interviews were conducted with 218 local contest participants and 198 participants from the statewide contest. Overall cessation impact (participation rate × abstinence) was 0.39 percent for the local contest and 0.09 percent for the statewide contest. Localized community contests offered in conjunction with statewide or national campaigns may represent cost-effective methods of reaching large numbers of smokers. (*Am J Public Health* 1990;80:601–603.)

Introduction

Contests to promote smoking cessation appear to represent cost-effective means of producing quit attempts in community settings.[1–3] Quit smoking contests have been offered on a number of occasions as part of the smoking intervention in the Minnesota Heart Health Program (MHHP), a 10-year research and demonstration project intended to reduce the prevalence of heart disease.[4,5]

Several smoking cessation contests have been timed to coincide with the Great American Smokeout conducted annually by the American Cancer Society (referred to as "D-Day" in Minnesota). The present study examined contest participation and outcome for samples of Twin Cities area residents in the 1984 Minnesota D-Day contest. Participants from one of the intervention communities (Bloomington) were compared with a random sample of those from other Minneapolis suburbs (not within the immediate Bloomington area). It was hypothesized that the overall impact of a contest, measured by participation and abstinence outcome, offered in conjunction with specific localized community recruitment and prizes would be greater than that of the statewide contest alone.

Method

Subjects were recruited for a statewide D-Day contest during the Fall of 1984. Recruitment began August 25, 1984

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optimistic to assume that similar results can be achieved elsewhere simply by adding a local promotion to a statewide contest.

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Smoking and Consistently High Use of Medical Care among Older HMO Members

DONALD K. FREEBORN, PHD, JOHN P. MULLOOLY, PHD, CLYDE R. POPE, PHD,
AND BENTSON H. MCFARLAND, MD, PHD

Abstract: Smoking behavior of consistently high and low users of medical care services were compared in two groups of older health maintenance organization (HMO) members continuously enrolled for five years and a subgroup who were continuously enrolled for 10 years. Smokers and former smokers, combined, were more likely than never-smokers to be consistently high users of ambulatory services (52 percent vs 34 percent in the five-year group, and 45 percent vs 30 percent in the 10-year group). (*Am J Public Health* 1990; 80:603–605.)

Introduction

Although the increased risk of morbidity and mortality associated with cigarette smoking is well established, the effect of smoking on the use of health services is less clear.[1-3] A number of studies have examined utilization patterns among the elderly,[4-7] but few have examined the relation between smoking and utilization. To our knowledge, no research has examined the relation between smoking and long-term ambulatory utilization patterns among persons 65 years of age and older.

This study seeks to determine: whether smoking behavior is an independent predictor of consistently high use of ambulatory medical services over periods of five and 10 years; and whether, when individuals are categorized as consistently high or low users of services, smokers (defined as current and former smokers) *within* the consistently high or consistently low user group use more medical services and resources than non-smokers (defined as persons who have never been smokers).

Methods

Data Sources

Study subjects were members of Kaiser Permanente (KP), Northwest Region, a large prepaid group practice HMO (health maintenance organization). Since January 1, 1967, medical record information has been routinely recorded for all medical contacts of a 5 percent sample of KP members.[8]

In 1970–71 the Center for Health Research at Kaiser Permanente conducted an extensive household interview survey of those subscriber units (families) in the 5 percent sample who were continuously enrolled in KP for the full two years of 1969 and 1970. The household survey was designed to obtain detailed individual and family demographic, social, economic, situational, behavioral, and attitudinal data, as well as health status data. The response rate was over 92 percent.[9] The survey and the outpatient utilization system for sample members age 65 and over (as of 1970–71) provide the data for this study.

Study Groups

Two study groups have been identified for this analysis:

- 312 sample members who were 65 years of age or older at the time of the survey, continuously enrolled in KP for the five-year period 1970–74, and alive at the end of 1975.
- A sub-group consisting of 203 sample members still alive at the end of 1980 who had been continuously enrolled for the 10-year period 1970–79.

Utilization Patterns

We determined the actual distribution of total ambulatory care contacts for each of the years 1970–74 for the first study group, and for the years 1970–79 for the second study group. Subjects were categorized as low, medium, or high users of services in each of the years according to the tertile into which their utilization fell.

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TABLE 1—Background Characteristics by Medical Care Service Utilization Pattern

Characteristics	Consistently Low Users	Consistently Mixed Users	Consistently High Users
Mean Age at Survey			
5-year group	70.7	71.5	71.2
10-year group	70.3	70.5	70.9
Percent Male			
5-year group	34.4	43.9	30.8
10-year group	27.7	34.1	18.2
Percent Married			
5-year group	54.7	63.3	65.4
10-year group	48.9	64.2	60.6
Percent High School Education or greater			
5-year group	29.7	38.1	42.3
10-year group	36.2	42.6	45.5
Income-Percent Less than $10,000			
5-year group	61.3	55.9	58.3
10-year group	65.2	55.1	60.7

TABLE 2—Smoking by Medical Care Service Utilization Pattern

	Consistently Low Users		Mixed Users		Consistently High Users		All Members	
	No.	%	No.	%	No.	%	No.	%
Group 1:1970–74								
Never smoked ($\bar{S}$)	42	65.6	127	66.1	25	48.1	194	62.9
Smoker, former smoker (S)	22	34.4	65	33.9	27	51.9	114	37.1
Total	64	100.0	192	100.0	52	100.0	308	100.0
High-Low Difference				17.5				
90% Confidence Interval				2.5–32.5				
Group 2:1970–79								
Never smoked ($\bar{S}$)	33	70.2	80	66.7	18	54.5	131	65.5
Smoker, former smoker (S)	14	29.8	40	33.3	15	45.5	69	34.5
Total	47	100.0	120	100.0	33	100.0	200	100.0
High-Low Difference				15.7				
90% Confidence Interval				−2.3–33.7				

Chi-squared test (Group 1) $p = 0.05$
Chi-squared test (Group 2) $p = 0.19$

For the first study group, those who were classified as high users in four or five of the years were defined as "consistently high users" (n = 52). Conversely, those classified as low users in four or five of the years were defined as "consistently low users" (n = 64). All others were defined as "mixed users" (n = 196).

For the second study group, those who were classified as high users in seven or more of the years were defined as "consistently high users" (n = 33). Those classified as low users in seven or more of the years were defined as "consistently low users" (n = 47). All others were defined as "mixed users" (n = 123).

Predictor and Control Variables

Both the predictor variable (smoking behavior) and the control variables (age, sex, marital status, and socioeconomic status) derive from the survey. Current and former smokers are combined in this study (and referred to as "smokers") because of the small number of cases.

In the five-year study group, 114 persons were included in the combined category of current and former smokers. Of this group, 67 percent were former smokers and 33 percent were current smokers. In the 10-year study group, the combined category included 35 persons (69 percent were former smokers and 31 percent were current smokers).

Analysis

To deal with potential confounding effects, we first examined the relation between the sociodemographic factors and the utilization patterns. None of the sociodemographic factors, individually or in combination, was associated with the utilization patterns (Table 1). The numbers, however, are small.

The analysis presented here is limited to the bivariate relationships between smoking behavior and patterns of use and, within categories of users, smoking behavior and resource use. The latter is defined as the number of office visits

TABLE 3—Medical Care Service Utilization by Utilization Pattern and Smoking Status

	Consistently Low Users			Mixed Users			Consistently High Users			All HMO Members		
Group 1; 1970-74 (N =)	Total (64)	S (22)	$\bar{S}$ (42)	Total (192)	S (65)	$\bar{S}$ (127)	Total (52)	S (27)	$\bar{S}$ (25)	Total (308)	S (114)	$\bar{S}$ (194)
Office visits per person per year	1.38	1.68	1.23	5.35	5.46	5.30	11.81	11.3	12.3	5.62	6.12	5.33
Hospital admissions per person per year	0.03	0.05	0.02	0.14	0.15	0.14	0.36	0.33	0.38	0.16	0.17	0.15
Average ambulatory costs per person per year+	$56	$72	$48	$208	$210	$207	$452	$440	$465	$218	$238	$206
Group 2: 1970–79	Total	S	$\bar{S}$	Total	S	$\bar{S}$	Total	S	$\bar{S}$	Total	S	$\bar{S}$
(N =)	(47)	(14)	(33)	(120)	(40)	(80)	(33)	(15)	(18)	(200)	(69)	(131)
Office visits per person per year	1.79	2.29	1.58	5.79	5.65	5.86	11.55	11.25	11.81	5.80	6.18	5.60
Hospital admissons per person per year	0.05	0.07	0.04	0.14	0.14	0.13	0.37	0.33	0.40	0.15	0.17	0.15
Average ambulatory costs per person per year+	$71	$90	$62	$225	$211	$232	$455	$414	$490	$227	$231	$225

+Costs are in terms of 1977 dollars

and hospital admissions per person per year, and the average per-person annual cost of ambulatory services.

Results

As shown in Table 2, for the five-year analysis, 52 percent of the high users compared to 34 percent of the low users were smokers (for the 10-year analysis, 46 percent vs 30 percent, respectively).

Table 3 shows that for both the five-year and the 10-year group, smokers had a higher rate of office visits and hospital admissions and greater average annual ambulatory costs per person than non-smokers. This pattern, however, does not obtain within the mixed user category.

Discussion

Elderly persons who were consistently high users of medical care were more likely than consistently low users to be smokers or former smokers; overall, smokers used more ambulatory care services than nonsmokers.

Limiting the study to long-term elderly HMO members may have understated the effects of smoking on the use of ambulatory medical services because of the differential survival rates of smokers and non-smokers. More smokers than nonsmokers may have been excluded from the study because of deaths. Nevertheless, the finding that, overall, non-smokers used less ambulatory services than smokers is consistent with the public health position that reductions in smoking will improve health status and reduce utilization.

Smoking cessation interventions can be successfully integrated into primary care practice and these interventions can be effective.[10] Long-term follow-up is needed to determine whether smoking cessation programs reduce subsequent utilization and the overall cost of medical care services.

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Routine Communication in Sexually Transmitted Disease Clinics: An Observational Study

Debra L. Roter, DrPH, Nancy Knowles, BA, Mark Somerfield, MS, and Julie Baldwin, BA

Abstract: Sixty STD patients were followed during the course of their clinic visit which averaged 2.75 professional contacts. Based on audiotape analysis, virtually all patients received a diagnosis; however, 25 percent of the patients did not receive any information about treatment or prevention, and only 57 percent of patients were given complete information. Patient recall of the information they were given averaged about 43 percent. Patients not seen by a physician during their clinic visit achieved higher recall rates than patients seen by a physician. (*Am J Public Health* 1990; 80:605–606.)

Introduction

Increasingly high rates of sexually transmitted diseases (STD) underscore the importance of Federal Quality Assurance Guidelines for STD Clinics.[1] According to guidelines, patient education should be routinely included in every patient visit and reinforced by every provider the patient sees. Validation of adherence to the Federal guidelines in STD clinics, however, has been infrequent.[2]

The current study was undertaken to observe directly provider-patient communication during routine STD clinic visits, to determine the extent to which diagnostic, treatment, and preventive information is given to patients as suggested by the Federal Guidelines, and to assess patient recall of this information.

Methods

Five health departments in the State of Maryland were asked to participate in a direct observation study of providers' practices in routine STD visits. One health department declined and another was dropped because of logistical difficulties. Observations were made in six clinics, representing three different county health departments between November 1985 and April 1986. Seventy-nine patients were enrolled in this study; this analysis is based on 60 patients for whom complete data were available. The patient sample was predominantly male (71 percent), and racially mixed (45

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Increase in Condom Sales following AIDS Education and Publicity, United States

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Abstract: Data from a national probability sample of drug stores show that condom sales rose from 240 million annually in 1986 to 299 million in 1988. The greatest increase occurred in 1987 after the Surgeon General's report on AIDS was released. Sales of latex condoms with spermicide rose 116 percent. Sales of other types of condoms increased less. These data suggest that Americans are using more condoms and probably more effective condoms in response to AIDS education. (Am J Public Health 1990; 80:607–608.)

Introduction

In his report on the acquired immunodeficiency syndrome (AIDS), the US Surgeon General advised persons whose sex partners may have been exposed to the AIDS virus to always use a condom during sexual intercourse.[1] An extensive public health campaign has since promoted condom use by persons who have sex outside of mutually monogamous relationships. Latex condoms, especially those containing spermicide, are believed to provide more protection than natural membrane condoms and have been the most highly recommended.[2]

Given the extensive public education effort and attendant publicity, we wondered if Americans were using more condoms. In a 1988 survey of urban teenage males, 58 percent reported using condoms during their last sexual encounter, compared to only 21 percent in 1979.[3] We analyzed a more objective measure, condom sales trends, to see how they have changed during the HIV (human immunodeficiency virus) epidemic.

Methods

Drug stores condom sales were estimated from bimonthly sales audits conducted by Nielson Marketing Research (Northbrook, IL). This national probability sample is stratified by size of store (four strata), geographic region (10 strata) and relative urbanization of the county where the store is located (four strata). Approximately 550 drug stores are audited, which compose 1 percent of drug stores and account for 2 percent of drug store sales. The same stores are audited every two months, making the system very sensitive to trends.

Condom sales in areas with a high incidence of AIDS were compared with sales in the remaining US. High incidence areas were defined as the television markets of the following metropolitan areas: Atlanta, Houston, Los Angeles, Miami-Fort Lauderdale, New York, San Francisco, and Washington, DC. These areas encompass one-fifth of the US population, one-third of all condom sales, and two-thirds of all AIDS cases. To reduce the tendency of seasonal variations in sales to obscure short-term trends, we compared sales for each two-month period with sales in the same period a year earlier.

To gauge condom publicity over time, we counted monthly condom citations using three indices: Newspaper Abstracts (University Microfilms International, Inc., Ann Arbor, MI), National Newspaper Index and Magazine Index (both produced by Information Access Company, Foster City, CA). All three showed the same pattern; only Newspaper Abstracts data are shown.

Results

Drug store condom sales grew slowly from 1984 to 1988 except for a 20 percent increase between 1986 and 1987 (Table 1). Sales of some styles grew more rapidly than others. Between 1986 and 1988, sales of *all* latex condoms increased 25 percent (226.1 to 283.4 million). The biggest percentage growth was in latex condoms with spermicide which increased 116 percent (23.2 to 50.1 million). Latex condoms without spermicide increased 15 percent (202.9 to 233.3 million). Natural membrane condom sales increased 7.8 percent (14.4 to 15.6 million).

Sales increased both in areas with a high incidence of AIDS and in the remaining US between 1986 and 1988 (Figure 1). Sales in the high incidence areas were growing throughout 17 of the 18 two-month periods. In contrast, sales were not growing in the remaining US until the beginning of 1987, and sales stopped increasing in July-August 1988. In both areas, condom sales grew rapidly throughout 1987 and early 1988 following the release of the Surgeon General's report in November 1986. Media attention to condoms also increased. Condoms were rarely mentioned before the report, but were increasingly cited in articles, editorials, and cartoons thereafter, reaching a peak in February 1987, when 182 items appeared in the 19 newspapers indexed by Newspaper Abstracts. (Forty percent of the February items concerned the controversy over whether condom advertisements were appropriate for television.) Throughout the remainder of 1987

TABLE 1—Annual Condom Sales in US Drug Stores, 1985–88, and Percent Change from Previous Year

	Sales in Millions (% change)			
	Latex			
Year	With Spermicide	Without Spermicide	Natural Membrane	Total
1985	22.4 (*)	201.4 (*)	13.9 (*)	273.7 (0.6)
1986	23.2 (3.2)	202.9 (0.7)	14.4 (3.9)	240.5 (1.2)
1987	35.0 (50.9)	237.1 (16.9)	17.3 (19.8)	289.4 (20.3)
1988	50.1 (43.2)	233.3 (−1.6)	15.6 (−10.0)	298.9 (3.3)

*data not available

Address reprint requests to Technical Information Services, Center for Prevention Services, Mailstop E-07, Centers for Disease Control, Atlanta, GA 30333. Dr. Moran is an Epidemic Intelligence Service Officer with the Clinical Research Branch, Division of STD/HIV Prevention, Center for Prevention Services, CDC. Dr. Peterman is Chief, Viral Studies Section, and Dr. Stone is Clinical Research Investigator, also at the Center for Prevention Services, CDC. Mr. Janes is Vice President, Product Management, Nielsen Marketing Research, Northbrook, IL. This paper, submitted to the Journal September 5, 1989, was revised and accepted for publication December 13, 1989.

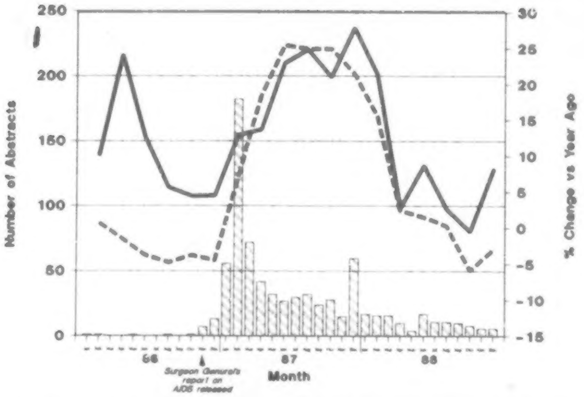

FIGURE 1—Annual Percentage Change in Condom Sales in Drug Stores in Seven Areas with High Incidence of AIDS (solid line) and remainder of United States (broken line) and Number of Newspaper Abstracts Mentioning Condoms (bars), 1986–88.
NOTE: Percent change in sales from the same period one year earlier are shown for two-month periods.

and 1988 media attention was greater than in 1986, but slowly diminished.

Discussion

The 20 percent increase in drug store condom sales in the year following the Surgeon General's report suggests that Americans responded to his message. In comparison, cigarette sales fell only 2.4 percent in the year following the Surgeon General's first report on smoking and health.[4]

The increase in sales was probably related to the recommendation that condoms be used to prevent HIV transmission. Sales of latex condoms, which were recommended for AIDS prevention, increased more than sales of natural membrane condoms. In addition, the greatest percent increase was in latex condoms with spermicide which cost more but may provide additional protection. The overall increase in drug store condom sales between 1984 and 1988 was 26 percent.

How much of the increase was for disease prevention? To answer this, we must estimate the proportion of condoms purchased primarily for contraception. According to studies of US women between 18 and 44 exposed to the risk of unintended pregnancy,[6] 14 percent of married women relied on condoms in 1982, making up 73 percent of the condom users in this population; in 1987, the proportion of married women using condoms was almost unchanged at 15 percent, suggesting that married women choosing condoms are mainly concerned with contraception, not disease prevention. In contrast, the percentage of unmarried women choosing condoms increased from 9 percent in 1982 to 16 percent in 1987, probably reflecting concerns about disease.[6] If all unmarried women choosing condoms do so for disease prevention, then there was a 78 percent increase in women using condoms for disease prevention.

How do sales compare with the number needed if all Americans who have intercourse outside of mutually monogamous relationships use condoms with every act of intercourse? Approximately 23 million Americans reported having more than one sex partner per year and 120 million reported a single sexual partner.[5] However, we do not know how many relationships are mutually monogamous or how often any of these people have intercourse. This lack of knowledge of Americans' sexual activity make it impossible to estimate the number of condoms needed to protect Americans from AIDS.

Other studies have shown increases in self-reported condom use among particular groups. However, behavior self-reports may be more affected by educational messages than behavior itself. Although this study has the advantage of using an objective measure of behavior, we have no evidence that the condoms purchased are actually used. Furthermore, the sales data are from drug stores only; condoms sold elsewhere or distributed free-of-charge are not included. Also, sales data give no information on who buys condoms or why.

This nationwide study of condom purchasing and other repeated surveys of sexual behavior show encouraging trends. But some populations are not being reached adequately. Despite the increase in sales and reported use of condoms, syphilis rates increased 25 percent between 1986 and 1987, with urban areas having the highest rates.[7] We need to learn how to increase the use of condoms among those who are at greatest risk of disease.

ACKNOWLEDGMENTS

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Car Design and Risk of Pedestrian Deaths

LEON S. ROBERTSON, PHD

Abstract: Fatal pedestrian injury rates by cars with relatively sharp front-corner designs were compared to such rates by cars of similar-size with relatively smooth front-corner designs. The relative risk of death by front-corner impact was 26 percent greater among the sharp-cornered cars. Pedestrian death rates from impact with other points on the cars and insurance claim frequencies among the studied cars were similar between the two sets of cars. (*Am J Public Health* 1990; 80:609–610.)

Introduction

In 1967, William Haddon, Jr., sent letters to 38 manufacturers of motor vehicles and their suppliers with pictures of designs that increased the severity of collisions with pedestrians.[1] In 1971, Haddon showed a committee of Congress the design of front ends of cars that resembled the design of ancient weapons—spears and battle axes.[2]

Despite Haddon's warnings, in the 1980s a number of large cars featured sharp protrusions as part of the design of their front corners. In impacts with pedestrians, these designs would be expected to increase the severity of injury because of their shearing effect.[3] The purpose of the research reported here was to investigate the extent to which such designs were associated with increased risk of pedestrian mortality in comparison with similar-sized cars that had relatively smooth front-corner designs.

Methods

A trade magazine with pictures of new cars introduced for sale in the United States during the model years 1980 through 1985 was examined to identify those that had sharp edges on their front corners.[4] Since all of the vehicles with sharp corners weighed more than 3,000 pounds, all cars from the same manufacturers that had relatively smooth front corners and weighed more than 3,000 pounds were chosen for comparison. The compared vehicles are listed in the Appendix.

Data on pedestrian fatalities in which the pedestrians were struck by the listed cars during the calendar years 1982–87 were obtained from the Fatal Accident Reporting System (FARS) computer files. These files contain data on virtually every fatal motor-vehicle injury in the US. The data include clock position of principal impact, eleven and one o'clock for front-corner impacts. Rates of pedestrian fatalities per vehicle years of use by principal point of impact were calculated using sales data[4] to estimate vehicle years of use.

To estimate vehicle use of vehicles sold during the calendar years of the study as accurately as possible, monthly sales data were used to calculate months of use during the year in which a given model was sold. The total years of use was thus the number of months from the sale in the first year divided by 12 plus the total number sold in the year times the number of years subsequently through 1987.

The rates of pedestrian deaths per 100,000 years of use from front-corner impacts and impacts at other points on the vehicles of the sharp-cornered relative to the smooth-cornered vehicles were compared. If differences in the pedestrian death rates in front-corner impacts of the two sets of vehicles were due to differences in driver factors, vehicle factors other than front-corner designs, or use in different environments, a similar difference in non-corner impacts should be found between the two sets of vehicles. There is no reason to believe that other risk factors would increase front-corner impacts but have no effect on pedestrian impacts at other points on the same vehicles.

Also, if the sharp-cornered vehicles were driven differently or had more defects that would increase their risk of collisions generally relative to the comparison vehicles, the insurance claim frequency for those vehicles should be higher. Therefore, the average insurance claim frequency of the two sets of vehicles was also compared using published data from insurance claims.

Results

The numbers of pedestrian deaths and deaths rates per years used in front-corner and other-point impacts of the cars with relatively sharp and relatively smooth front corners are presented in Table 1. The death rate of the relatively sharp-cornered cars in front-cornered impacts was 1.26 times that of the relatively smooth-cornered cars (95% CI = 1.05, 1.50). Thus the sharp-cornered vehicles had an average of 26 percent greater risk of pedestrian deaths in front-cornered impacts. In impacts with other points on the vehicles, the sharp-cornered cars had a relative risk of 0.98 compared to the smooth-cornered cars (95% CI = 0.86, 1.11), that is, virtually the same rates in non-cornered cars.

TABLE 1—Relative Risk of Fatal Pedestrian Injury By Front-corner Design and Point of Impact—Large US Cars, 1982–87

Front-Corner Design	Deaths			Death Rate		Relative Risk	
	Front Corner	Other	100,000 Years Used	Front Corner	Other	Front Corner	Other
Smooth	589	1403	641.9	0.91	2.19	1.00	1.00
Sharp	160	296	138.8	1.15	2.13	1.26	0.98

95 percent confidence intervals (log odds method) (Front corner 1.05, 1.50) (Other 0.86, 1.11).

Address reprint requests to Leon S. Robertson, PhD, Nanlee Research, 2 Montgomery Parkway, Branford, CT 06405. This paper, submitted to the Journal August 31, 1989, was revised and accepted for publication October 27, 1989.

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The insurance claim data were relative claim frequencies standardized so that the all-car average equals 100.[6] The smooth-cornered cars in this study had an average relative claim frequency of 86. The claim frequencies of the sharp-cornered cars had an average of 87, little different from the smooth-cornered cars' average.

Discussion

As physics and biomechanics predict, moving objects with sharp points do more damage when they strike people than do smooth objects.[3] The 26 percent average greater risk of fatal pedestrian injury from sharp-cornered compared to smooth-cornered cars is consistent with that knowledge. The rate of fatal non-corner impacts, comparing the same vehicles, were similar, suggesting that the major difference among the vehicles in risk to pedestrians is the sharpness of their front corners. The average insurance claim frequency of the two sets of vehicles were nearly the same. Therefore, it is unlikely that driver, environmental, or other vehicle factors accounted for the differences in pedestrian death rates from front-corner impacts.

Vehicle manufacturers often cite costs as the reason for refusing to design and build safer vehicles. Yet the sharp corners on the vehicles in this study require more material than would smooth corners.

The lack of difference in rates of non-corner impacts does not necessarily mean that other points on the vehicles were optimally designed. They were just not different, on average, among these two sets of vehicles. Low bumpers and hoods in combination with smooth, sloped hoods and energy absorbing material in the bumpers and windshield rims are the most efficacious designs.[7]

APPENDIX

Vehicles Studied

Sharp Corners

- Buick Riviera
- Cadillac Eldorado
- Chrysler Cordoba
- Chrysler New Yorker
- Chrysler Fifth Avenue
- Chrysler Imperial
- Dodge Diplomat
- All Lincolns except the 1983–85 Continental and Mark VII
- Mercury Marquis
- Plymouth Fury

Smooth Corners

- Buick Electra
- Buick LeSabre
- Buick Regal
- Cadillac Deville
- Cadillac Fleetwood
- Cadillac Seville
- Chevrolet Caprice
- Chevrolet Impala
- Chevrolet Malibu
- Chevrolet Monte Carlo
- Ford LTD
- Ford Crown Victoria
- 1983-1985 Lincoln Continental and Mark VII
- 1980–1982 Mercury XR-7
- 1983–1985 Mercury Cougar
- Oldsmobile 88 and 98
- Oldsmobile Cutless (except Ciera)
- Pontiac Bonneville
- Pontiac Gran Prix

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APHA Epidemiology Section Invites Abstracts for Poster Session

The Epidemiology Section of APHA will again be sponsoring a "Late-Breaker" poster session at the upcoming annual meeting. The session will take place on Wednesday, October 3, 1990, 12:30–2:00 p.m. Work completed in 1990 is eligible for consideration. Abstracts must be *received* by September 7 at the address given below. Students and recent graduates are particularly encouraged to submit abstracts. Submit an abstract of less than 200 words (any format) and a return envelope to:

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Males' Use of Public Health Department Family Planning Services

MICHAEL GLASSER, PHD

Abstract: Information was obtained from over 200 males visiting a family planning service either to obtain free-of-charge condoms or incidental to accompanying their female partner. The former group was younger and more likely to reside near the health department; over 40 percent of this group used the health department as sole condom source. Both groups indicated willingness to share in costs for male services. Future programs need to more effectively plan public health family planning for these males. (*Am J Public Health* 1990; 80:611–612.)

Introduction

A pilot study, summarized in detail elsewhere, presented the responses of 70 male clients to a brief questionnaire concerning sexual activity, contraceptive use, and health department visits. Results indicated the health department was serving a substantial population of males and was a central and often exclusive source of condoms. Methodologically, under conditions of anonymity, males were willing to answer questions regarding family planning and sexual activity.

The present research was conducted at the Winnebago County Public Health Department in Rockford, Illinois (population 139,712), with 279,514 living in the Standard Metropolitan Statistical Area. The family planning service of the health department has focused in the past nearly exclusively on female clients, consistent with national trends. Danielson, *et al*, point out that formulas for fund allocation are derived primarily from the number of women in need of services as well as actual number served; men are considered only more peripherally, if at all.

In the present case, a condom distribution service for males was made possible through health department commitment to providing total services related to birth control needs. The family planning clinic set up a distribution desk for condoms where males could request the contraceptive at specific times. Contraceptive availability was advertised in brochures and in the newspaper. Males had to be at least 13 years old to receive condoms. They also received literature regarding family planning and contraceptive use. This study was designed to further provide information on males' use of these services.

Method

Results of the pilot study revealed that there were two different "groups" of males at health department family planning: males using the free-of-charge condom service (Group C) and males attending the family planning visit of their female partner (Group P). We obtained information from both groups.

Sampling focused on units of time broken down into three parts for each day that the condom distribution service and family planning service were open over a six-month data collection period. Nearly 30 percent of time periods were sampled. During these periods, all males requesting condoms or visiting with their partner were asked to complete a questionnaire or consent to an interview. At the time of this study, the Group P males were specifically visiting with their partners and did not request condoms from the distribution service during that visit. While some males visited family planning more than one time during the study period, they completed only one questionnaire or interview.

Results

A total of 207 males responded (74 percent from Group C and 78 percent from Group P). For Group C, there were 157 separate contacts during the study period, for Group P 144 encounters; altogether, an estimated 2,007 (±46 at 95% confidence intervals = 1961, 2053) contacts with males in one year.

Group C males using the condom service were younger than Group P males, less likely married, had a lower educational level, and lived nearer the health department. They reported a younger age of first intercourse—12.7 compared to 14.4 years. Over 47 percent of Group C had been sexually active by 12 years old compared to 26.2 percent of Group P. At first sexual encounter, nearly 60 percent of all males reported *not* using contraception.

When asked who has primary contraceptive responsibility, Group P males were more likely to indicate both the man and woman and relied less on condoms and more on other measures of birth control, such as the pill. Nearly one-half of Group C males reported condom use for two or more years but 15 percent used condoms for three months or less; over 45 percent indicated always using condoms while nearly 30 percent reported frequency of use as "usual."

Nearly two-thirds of Group C males were repeat visitors to the service, and over 40 percent had visited within one month (Table 1). Almost 42 percent reported the service as sole source for condoms. Main reasons for use were the service was free, friend's suggestion, and no questions asked. Nearly 70 percent visited the service with a friend.

For Group P males, about 60 percent were accompanying their partner to family planning for the first time. The males were there primarily to provide emotional support, give transportation, and be a part of the family planning process. Nearly one-fourth of Group P males had used the condom distribution service.

When asked if they would buy condoms from the service if necessary or share in expenses for services, 72 percent of Group C and 77 percent of Group P males replied yes (Table 2). Both groups reported satisfaction with health department services. The Group C males were a little more likely to indicate interest in additional services, including medical examinations, sexually transmitted disease (STD) programs, and support for adolescent fathers (66 vs 55 percent).

Discussion

The health department provides access to birth control

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TABLE 1—Characteristics in Use of the Health Department Family Planning Condom Distribution Service by 100 Males Visiting only to Receive Condoms

First Visit to Condom Service	
Yes	33
No	65
No Answer	2
Length of Time Since Last Visit	
Less Than 2 Weeks Ago	10
2 to 4 Weeks Ago	18
1 to 2 Months Ago	18
More Than 2 Months Ago	19
First Visit	33
No Answer	2
Family Planning as Only Source for Condoms	
Yes	41
No	57
No Answer	2
Reason(s) for Using Condom Service	
Free of Charge	55
No Questions Asked	32
Male Friend Suggested	21
Female Friend Suggested	25
Sent by Health Personnel	3
Sent by Adult	9
Location was Convenient	14
Other	10
In Company of Another Person	
Yes	68
No	30
No Answer	2

TABLE 2—Evaluation of Health Department Services by Group C and Group P Males

	% (n = 100)	% (n = 107)
Willing to Purchase Condoms or Share in Costs of Family Planning	**Group C**	**Group P**
Yes	72	77
No	28	23
No Answer	(2)*	(4)
Satisfaction Level		
Very Satisfied	51	50
Satisfied	26	22
No Opinion	20	25
Somewhat Dissatisfied	2	3
Very Dissatisfied	1	0
	(3)	(3)
Health Department Should Offer More Services for Men		
Yes	66	55
Not Sure	30	24
No	5	21
No Answer	(3)	(2)

*( ) indicates number of missing cases.

both for men wanting condoms and those whose partners are seeking family planning. For males using the condom distribution service, nearly one-half reported exclusive use of condoms, two-thirds made more than one visit to the health department, and 40 percent viewed the health department as sole condom source.

Many Group C males were "repeat" service users. Of the 1,047 projected visits, one-third (346) were first visits. Although it is not clear how many different males are represented in the 701 remaining visits, it is apparent that a large proportion of males could be classified as regular clients.

Group P males comprise a population "in place" to benefit from public health services. They expressed interest in obtaining general as well as family planning services and seem relatively "untapped" as a target of programming.

Study limitations were that data were collected at one time rather than longitudinally and the focus was on one facility. Still, it was evident that there was a high frequency of male contacts, if not direct service utilization. In relation to Title X and to health care in general, it appears there is a large population that could be brought more formally under the "umbrella" of family planning.

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